A Critical Review of Housing Markets

By Mr. Pin-Te Lin

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Signed Statement of Originality

The work presented in this thesis, is to the best of my knowledge my own work, except as acknowledged in the text. The material has not been submitted, either in whole or in part, for a degree at this or any other university.

Pin-Te Lin

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Abstract

This PhD dissertation, composed of three studies, provides a critical review of the determinants of price changes in the housing market. In the first paper, I address the counterintuitive negative risk-return relationship in the housing market, previously found in the literature. I show that the negative relationship in the US can be resolved by considering the modelling differences between aggregate and cross-sectional conditions in standard asset pricing theory. The result has implications for the application of standard finance theory towards housing markets.

In the second paper, I revisit the commonly-held beliefs that the nature of housing markets is mainly local and that local time-invariant amenities are crucial for understanding housing price dynamics. Based on the empirical evidence of the US housing markets, I show that the documented evidence in support of these arguments can be an artefact of sample size in the time series dimension of a panel data analysis. The result provides implications for empirical and theoretical research regarding the assumptions of housing market dynamics.

The third paper presents a historical review of the long-run relationship between macroeconomic factors and housing markets from 1871 to 2012. I find that century-long evidence across 10 countries has consistently favoured the role of inflation hedging in residential real estate, yet its significance has decreased in the recent regime of inflation targeting from 1990 to 2012. During the latter period, much of the variation in housing markets is linked not to inflation risk, but rather to changes in real income. The result adds a new dimension to understanding how inflation hedging benefits can change under different momentary environments.

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Paper 1. Back to the Beginning: Risk-Return Relationship in Housing Markets

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Abstract

While mainstream theory predicts a positive relationship between risk and return, empirical evidence shows a puzzling negative risk-return trade-off in some housing markets. This study resolves the source of the contradiction by concentrating on the modelling differences in conditional risk between the aggregate and cross-sectional conditions in Merton's (1973) Intertemporal Capital Asset Pricing Model (ICAPM). Based on the corresponding market clearing conditions, I show that the US housing market displays a significantly positive risk-return trade-off at both regional and national levels. I therefore demonstrate how standard finance theory can resolve the risk-return puzzle for housing.

*This research is the first paper of my PhD dissertation at the Australian National University. I am grateful to my supervisors for their support and guidance: Dr Dean Katselas (Chair) and Dr Anna von Reibnitz. I also thank conference participants at the 2016 American Real Estate Society Annual Meeting, 2016 Financial Research Network Annual Conference and 2016 Pacific Rim Real Estate Society Annual Conference for helpful comments and suggestions. Please direct correspondence to Mr. Pin-Te Lin, Research School of Finance, Actuarial Studies & Statistics, Australian National University, Canberra, ACT, 2601, Australia. E-mail: pinte.lin@anu.edu.au.

1. Introduction

The trade-off between risk and return is a fundamental concept in finance. Among all capital assets, the risk-return relationship in housing is particularly worthy of attention, since housing constitutes two-thirds of the typical American household's financial portfolio (Bayer, Ellickson, & Ellickson, 2010). Further, changes in housing wealth are known to have significant effects on consumption and investment decisions (Lovenheim & Mumford, 2013). Yet, despite its significance, little is known about how housing returns vary with price risk. While standard finance theory generally predicts a positive relationship between the market risk premium and conditional risk for risk-averse agents, a large body of empirical studies shows a negative relationship in housing markets (e.g., Dolde & Tirtiroglu, 1997; Han, 2013; Karoglou, Morley, & Thomas, 2013; Lin & Fuerst, 2014; Miles, 2008, 2011; Morley & Thomas, 2011, 2016; Willcocks, 2010; Zhou, 2016). The conflicting result is problematic since it implies that most housing investors can be risk-seeking.

This paper aims to resolve the empirical contradictions through the lens of aggregate and cross-sectional conditions in Merton's (1973) Intertemporal Capital Asset Pricing Model (ICAPM). I advocate that housing markets follow the intertemporal mean-variance efficiency of a theoretically positive risk-return trade-off. I argue that the documented contradictions in the housing literature are potential artefacts of unsatisfactory estimation from which researchers are compelled to draw inferences. This argument is supported by the fact that most equity literature highlights that the direction of risk-return trade-off can be highly sensitive to the way conditional risk is modelled (e.g., Frazier & Liu, 2016; Ghysels, Santa-Clara, & Valkanov, 2005; Guo & Whitelaw, 2006; Harvey, 2001; Lundblad, 2007; Nyberg, 2012; Pastor, Sinha, & Swaminathan, 2008; Rossi & Timmermann, 2015; Scruggs, 1998; Yu & Yuan, 2011). Consistent with this, I attribute the source of the contradiction to the modelling of conditional risk for examining the aggregate and cross-sectional conditions in Merton's (1973) ICAPM.

Through the families of Generalised Autoregressive Conditionally Heteroskedastic in the Mean (GARCH-M) models, the housing literature (e.g., Dolde & Tirtiroglu, 1997; Han, 2013; Karoglou et al., 2013; Lin & Fuerst, 2014; Miles, 2008, 2011; Morley & Thomas, 2011, 2016; Willcocks, 2010; Zhou, 2016) shows a significantly positive risk-return trade-off in some regional markets but a significantly negative trade-off in others. Underlying this contradiction is the presumption that GARCH-M analysis is a reasonable approach for examining the theoretical risk-return relationship in regional markets. The empirical argument motivating this paper is that the documented sample sizes (from 48 to 264 observations) used in previous work are insufficient for GARCH-M analysis, based on the criteria of Lundblad (2007).[1] In this context, I argue that the national Case-Shiller home price index contains the large number of time series observations that could satisfy this requirement. The theoretical argument motivating this paper is that regional markets are sub-markets within a nation; therefore, the risk measurement for regional markets should be conditional *covariance* rather than conditional *variance* according to the cross-sectional condition in Merton (1973). From this perspective, I argue that the theoretical market clearing condition for an aggregate market is preferably investigated at the national level with the application of GARCH-M models.

Putting the empirical and theoretical motivations together suggests one implication: documented contradictions throughout the literature are potentially driven by unsatisfactory risk measurements adopted per Merton's (1973) ICAPM. I argue that

[1] See Table 1 for details in previous work.

most claimed research findings in the housing literature against the theorised relationship in standard finance theory are questionable, unless more reasonable risk measurements are considered. The present paper is an exploration of this issue, with two sets of goals. First, based on the aggregate condition, I investigate whether the national housing market possesses a significantly positive risk-return trade-off between the market risk premium and conditional variance. Second, based on the cross-sectional condition, I examine whether the regional housing market displays a significantly positive risk-return trade-off between the market risk premium and conditional covariance. By addressing these questions, I shed new light on how to resolve the risk-return puzzle in the housing literature.

My empirical analysis begins by examining the aggregate condition in the national housing market of the US, with the application of the Case-Shiller home price index. I find a significantly positive risk-return trade-off between the market risk premium and conditional volatility through the standard GARCH-M models. My results are robust irrespective of the model specification (GARCH-M, Threshold GARCH-M, Exponential GARCH-M and Component GARCH-M) and error distribution (normal, student's t and generalised). The evidence of a significantly positive intertemporal relationship between the mean and variance in the aggregate housing market supports the fundamental prediction of the underlying asset pricing works of Merton (1973), Merton (1980) and Campbell (1993). Given the mixed evidence from standard GARCH-M analysis in the literature on equities, the ability of GARCH-M to detect a significantly and consistently positive risk-return trade-off in the US housing market has implications for researchers in considering whether the housing market might be a better natural setting for the ICAPM.

I next assess the cross-sectional condition for regional markets in the US. Since identifying the true risk-return relationship by the families of GARCH models has been acknowledged to be uninformative for a small sample size (e.g., Lundblad, 2007), I follow Han (2013) by pooling the conditional risk of each regional market into a panel data regression. Though my approach is similar to Han (2013), I echo Bayer et al. (2010) by highlighting that standard finance theory implies that what matters for the risk premium of housing assets is their exposure to systematic risk. Therefore, the adopted risk measure is the conditional covariance predicted by bi-variate GARCH modelling rather than the total risk generated by univariate GARCH modelling. In contrast to Han (2013), my empirical analysis of the cross-sectional condition in Merton's (1973) ICAPM uncovers a significantly positive risk-return trade-off across regional markets in an intertemporal framework.

The conditional analysis in this study can be viewed as an expansion of unconditional works in Beracha and Skiba (2013), Cannon, Miller and Pandher (2006) and Case, Cotter and Gabriel (2011). By treating housing as a single-period investment with static systematic risk, these works consistently show that housing markets of higher systematic risk are rewarded with higher returns. My multi-period conditional approach further shows that the principle of a positive risk-return trade-off can hold not only across sub-markets but also *over time*. The significance of this analysis has implications for investment strategy in housing, especially for investors in multi-family real estate investment trusts (REITs) and single-family housing investment funds, as highlighted in Cotter, Gabriel and Roll (2015).

This paper also combines and extends two separate strands of research in housing: one stream on conditional analysis (e.g., Dolde & Tirtiroglu, 1997; Han 2013; Karoglou et al., 2013; Lin & Fuerst, 2014; Miles, 2008, 2011; Morley & Thomas, 2011, 2016; Willcocks, 2010; Zhou, 2016) and the other stream on unconditional analysis (e.g., Beracha & Skiba, 2013; Cannon et al., 2006; Case et al., 2011). Fundamentally, I illustrate that the empirical analysis of the two schools of thought can be theoretically equivalent to a single-factor model. The difference between the two lines of investigation reflects the underlying assumption of whether the housing is a single or multi-period investment. The adopted assumption further leads to the choice between conditional and unconditional analysis. While one strand of the literature on unconditional analysis has reached consensus regarding a positive risk-return trade-off, the other strand on conditional analysis remains inconclusive. In this context, this paper's contribution is to identify the source of the contradiction in conditional analysis, providing a simple and intuitive empirical analysis of intertemporal mean-variance efficiency in housing.

More broadly, this paper engages with a substantial literature that investigates the risk-return trade-off for financial assets. Though a considerable body of literature concentrates on equities, attention on housing has been limited (Han, 2013). In fact, Roll's (1977) critique argues that the definition of the market portfolio is theoretically and empirically elusive. A comprehensive market portfolio should in principle include not just traded financial assets, but also consumer durables, real estate and human capital (Fama & French, 2004). As emphasised by Bayer et al. (2010), the relative omission of the investment side of housing analysis is surprising, since housing represents approximately two-thirds of the average American household's financial portfolio. Hence, this paper considers housing assets from a standard asset-pricing perspective, employing Merton's (1973) ICAPM to relate housing returns to their price risk in a dynamic framework. Through a complete examination of the aggregate and cross-sectional conditions in Merton (1973), this research on housing complements two lines of investigation in the equity literature: one on aggregate analysis (e.g., Frazier & Liu, 2016; Ghysels, Guerin, & Marcellino, 2014; Ghysels et al., 2005; Guo & Whitelaw, 2006; Jiang & Lee, 2014; Lundblad, 2007; Nyberg, 2012; Rossi & Timmermann, 2015; Salvador, Floros, & Arago, 2014; Scruggs, 1998; Wang, 2005; Yu & Yuan, 2011) and the other on cross-sectional analysis (e.g., Bali, 2008; Bali & Engle, 2010).

Finally, this study is closely related to the work of Han (2013), in that I also aim to resolve the risk-return puzzle in housing markets. However, my approach is markedly different. Presuming that the GARCH-M model is a satisfactory methodology for examining the theoretical risk-return relationship in regional markets, Han (2013) views the contradictory signs of intertemporal risk-return relationship as evidence against the postulated relationship in Merton's (1973) ICAPM. The different approach of this paper is to highlight that the conflicting risk-return relationship in housing markets does not relate to a failure of standard asset pricing theory, but rather to an assumption issue. Overall, my empirical evidence using aggregate and cross-sectional analyses consistently indicates that housing markets obey the intertemporal mean-variance efficiency. Therefore, I conclude that standard finance theory can resolve the argued risk-return puzzle in the housing literature.

2. Research Motivation

This section begins by surveying the related housing literature. I document that the empirical contradictions of the mean-variance relationship in housing could result from the measurement of conditional risk using Merton's (1973) ICAPM in both empirical and theoretical settings. By addressing this issue, I raise the possibility that standard finance theory can explain the risk-return relationship in housing markets.

2.1 Literature Review

To better understand the financial characteristics of housing markets, there has been a resurgence of research addressing the link between risk and return in this asset class. However, empirical results have been contradictory, as summarised in Table 1. Previous work based on different versions of univariate GARCH-M models fails to agree about the relationship between risk and return across regional markets.

For example, at the state level, Miles (2008) finds a significantly positive risk-return relationship in Georgia but a significantly negative relationship in Nebraska. At the city level, Karoglou et al. (2013) detect a significantly positive risk-return relationship in Portland but a significantly negative relationship in Chicago. Unlike these studies, however, Han (2013) goes one step further and reconciles conflicting findings by considering the dual financial-asset and consumption-hedge role of housing. This paper complements Han's (2013) solution to the puzzle by considering the modelling differences between aggregate and cross-sectional conditions in Merton's (1973) ICAPM.

Table 1. Summary of the risk-return literature

Paper	US market	Sample period (frequency)	Numbers of return observations	Method	R-squared	Conclusion
Dolde and Tirtiroglu (1997) ♦	Regional	Connecticut: 1982:Q1–1994:Q1 San Francisco: 1970:Q1–1988:Q1 (quarterly)	Connecticut: 48 San Francisco: 72	GARCH-M	Not reported	Significantly positive for San Francisco/ inconclusive for Connecticut
Miles (2008) ♣	Regional	1979:Q1–2006:Q2 (quarterly)	109	GARCH-M	Not reported	Insignificant for some states/ significantly negative for some states/ significantly positive for some states
Karoglou et al. (2013) †	Regional	1987:M1–2009:M1 (monthly)	264	Component GARCH-M	Not reported	Insignificant for some cities/ significantly negative for some cities/ significantly positive for some cities
Han (2013) ※	Regional	1980:Q1–2007:Q4 (quarterly)	111	GARCH-M/ Two-step estimation	Not reported	Significantly negative for some cities/ significantly positive for some cities

Notes: There is a substantial literature in intertemporal mean-variance analysis of housing. For representative work, I select the articles from the generally accepted top three real estate finance journals: † Journal of Real Estate Finance and Economics, ♣ Journal of Real Estate Research and ♦ Real Estate Economics. I also include one article from one of the top three finance journals: ※ Review of Financial Studies.

2.2 Theoretical Motivation

Merton's (1973) ICAPM generates two market clearing conditions for the risk-return relationship: aggregate and cross-sectional. In this paper, I argue that risk-return investigation at the regional level is consistent with the theoretically cross-sectional condition, while investigation at the national level is consistent with the theoretically aggregate condition. Assuming the existence of risk-averse representative agents, the cross-sectional condition of Merton (1973) indicates that the excess return of an asset or a portfolio *i* ($R_{i,t+1}$) is a linear function of its conditional covariance with the aggregate market ($\sigma_{im,t+1}$) and covariance with investment opportunities ($\sigma_{is,t+1}$):

$$E_t[\,R_{i,t+1}] = \left(\frac{-J_{ww}W}{J_w}\right)\sigma_{im,t+1|t} + \left(\frac{-J_{ws}W}{J_w}\right)\sigma_{is,t+1|t}, \tag{1}$$

where *W(t)* is wealth, *S(t)* is a variable describing the state of investment opportunities in the economy and *J (W,S,t)* represents the investor's indirect utility function. Subscripts of *J* denote partial derivatives. $(-J_{ww}W/J_w)$ is the related risk aversion and is assumed to be positive, indicating a positive risk-return trade-off for risk-averse agents.

I further form an aggregate market condition by summing up Equation (1) from each of the regional markets and multiplying by its corresponding market weights (w_i):

$$E_t\left[\sum_1^n w_i\,(R_{i,t+1})\right] = \sum_1^n w_i\left[\left(\frac{-J_{ww}W}{J_w}\right)\sigma_{im,t+1|t} + \left(\frac{-J_{ws}W}{J_w}\right)\sigma_{is,t+1|t}\right], \tag{2}$$

where $R_{m,t+1} \equiv \sum_1^n w_i\left(R_{i,t+1}\right)$, $\sigma^2_{m,t+1|t} \equiv \sum_1^n w_i\sigma_{im,t+1|t}$ and $\sigma_{ms,t+1|t} \equiv \sum_1^n w_i\sigma_{is,t+1|t}$.

$$E_t\left[R_{m,t+1}\right] = \left(\frac{-J_{WW}W}{J_W}\right)\sigma^2_{m,t+1|t} + \left(\frac{-J_{WS}W}{J_W}\right)\sigma_{ms,t+1|t}. \tag{3}$$

The derived Equation (3) shows that the conditional excess return on the aggregate market is a linear function of its conditional variance and covariance with investment opportunities (the hedging component). If we assume that a regional housing market is part of a national market, then the progression from Equations (1) to (3) quantitatively proves that the cross-sectional condition is the corresponding market clearing condition for local markets.

The contradictory results in several housing studies, including Dolde and Tirtiroglu (1997), Miles (2008, 2011), Willcocks (2010), Morley and Thomas (2011), Han (2013), Karoglou et al. (2013), Lin and Fuerst (2014), Morley and Thomas (2016) and Zhou (2016), reflect the assumption that GARCH-M analysis is a reasonable approach for examining the theoretically intertemporal risk-return relationship in regional markets. However, this assumption is at odds with standard finance theory, since what accounts for the risk premium of regional markets should be their exposure to systematic risk rather than total risk.

Further, as is well documented in the equity literature (e.g., Bali & Engle, 2010; Frazier & Liu, 2016; Ghysels et al., 2005, 2014; Guo & Whitelaw, 2006; Jiang & Lee, 2014; Lundblad, 2007; Nyberg, 2012; Pastor et al., 2008; Rossi & Timmermann, 2015; Salvador et al., 2014; Scruggs, 1998; Yu & Yuan, 2011), the aggregate condition between market risk premium and conditional variance (i.e., total risk) in Merton (1973) is a unique circumstance, typically investigated at the national level. If regional housing markets are part of the aggregate wealth portfolio, it follows that systematic risk should be the major pricing factor for regional markets, according to the cross-sectional condition in Merton (1973).

2.3 Empirical Motivation

The literature recognises that GARCH models based on small samples will produce unreliable estimates for conditional volatility (e.g., Guo, Wang, & Yang, 2013; Hill & Prokhorov, 2016; Hwang & Pereira, 2006; Iglesias & Phillips, 2011; Lundblad, 2007; Willcocks, 2009). As highlighted in Hwang and Pereira (2006), one needs at least 500 observations to estimate using a standard GARCH(1,1) model. Surveying related work, as shown in Table 1, I have found that previous work consistently applies the families of GARCH-M models to regional markets with small sample sizes—from 48 to 264 observations. A more direct study by Lundblad (2007) further shows that GARCH-M analysis based on a small sample can easily result in an insignificant or even a significantly negative risk-return relationship even if the true risk-return trade-off is positive. Using a small sample size of 500 observations in a standard GARCH-M model, Lundblad's (2007) simulation shows that 19% of the mean-variance parameters can fall below zero. Therefore, it is no surprise that previous work has been contradictory.

Table 2. GARCH-M results in the US housing market

	Sub-sample (Obs:81)			Full sample (Obs:162)		
Error term distribution	(1) Normal	(2) Student's t	(3) Generalised	(1) Normal	(2) Student's t	(3) Generalised
c	-0.004***	-0.004***	-0.004***	-0.001	-0.001	-0.002*
	(0.002)	(0.002)	(0.002)	(0.001)	(0.001)	(0.001)
α_M	-3.176	-3.114	-3.817	-1.036	-1.035	0.129
	(12.621)	(12.585)	(12.689)	(7.249)	(7.254)	(7.967)
[Relative strength to the sub-sample]				[More positive]	[More positive]	[More positive]
δ_0	1.7×10^{-6}	1.8×10^{-6}	1.5×10^{-6}	3.7×10^{-6}	3.7×10^{-6}	2.6×10^{-6}
	(4.3×10^{-6})	(4.6×10^{-6})	(3.7×10^{-6})	(2.7×10^{-6})	(2.7×10^{-6})	(2.3×10^{-6})
δ_1	0.161*	0.164*	0.157**	0.340***	0.340***	0.264***
	(0.086)	(0.093)	(0.073)	(0.129)	(0.130)	(0.084)
δ_2	0.805***	0.803***	0.809***	0.677***	0.677***	0.740***
	(0.079)	(0.087)	(0.069)	(0.081)	(0.082)	(0.062)
Adjusted R-squared	-0.035	-0.034	-0.041	-0.010	-0.010	-0.020
[Relative explanatory power to the sub-sample]				[Higher]	[Higher]	[Higher]

Notes: The table presents the coefficient estimates for a standard GARCH-M (1,1) model with the mean equation: $r_{m,t} - r_{f,t} = c + \alpha_M \sigma_{M,t}^2 + \varepsilon_t$ and variance equation: $\sigma_{M,t}^2 = \delta_0 + \delta_1 \varepsilon_{t-1}^2 + \delta_2 \sigma_{M,t-1}^2$. Numbers in parentheses denote the corresponding standard errors for the coefficients. ***, ** and * denote significance levels at 1%, 5% and 10%, respectively. Market returns are calculated as $r_t = ln(P_t/P_{t-1})$ where P_t is the national housing price index in the US from the Federal Housing Finance Agency (FHFA) over the period 1975:Q1 to 2015:Q3. The proxy for the risk-free rate is one month treasury bill rate from the Center for Research in Security Prices (CRSP). The sub-sample is the first half of the sample period.

To illustrate the concept of the small sample size problem, I apply a GARCH-M model to the national housing market of the US. Though my application of a GARCH-M model at the national level maintains the theoretical consistency of an aggregate condition, empirical evidence of the risk-return relationship in Table 2 remains insignificant. Nonetheless, carefully comparing the results of the sub-sample to the full sample, I find that the direction of the risk-return relationship (i.e., the coefficient estimates for α_M) and explanatory power (i.e., the adjusted R-squared) become positive and increase, respectively, upon moving to the larger sample. These results are consistent with the predictions in Lundblad (2007). As yet, while I am agnostic as to whether the prediction of standard finance theory can hold in the national housing market, the empirical evidence does imply that the weak empirical risk-return relationship may be a detection problem. As I next show in the main empirical analysis, one can successfully identify a significantly positive risk-return relationship in the US housing market when a large amount of time series data are used through the national Case-Shiller home price index.

2.4 Argument

Combining the empirical and theoretical motivations leads to one implication: Merton's (1973) ICAPM could be alive and well in housing markets. Though Han (2013) treats the empirical contradictions as evidence against the theorised relationship, I argue that the conflicting results stem from a variety of empirical inspections that employ extensively divergent specification and estimation methods. To demonstrate, several studies in Miles (2008), Karoglou et al. (2013) and Han (2013) cite that Dolde and Tirtiroglu (1997) find a negative risk-return relationship in Connecticut. However, Dolde and Tirtiroglu's (1997) empirical evidence for the negative relationship in Connecticut is only significant in one out of the four versions of GARCH-M models. Therefore, unlike the majority of the housing literature, this study takes a new look at Merton's (1973) ICAPM towards housing markets, attributing conflicting results to risk measurements rather than regarding them as absolute evidence against the postulated relationship.

In parallel, the extensive literature on equity also advocates my proposed argument (e.g., Frazier & Liu, 2016; Ghysels et al., 2005, 2014; Guo & Whitelaw, 2006; Harvey, 2001; Jiang & Lee, 2014; Lundblad, 2007; Nyberg, 2012; Rossi & Timmermann, 2015; Salvador et al., 2014; Scruggs, 1998; Wang, 2005; Yu & Yuan, 2011). For example, in the context of GARCH models, one of the pioneering studies by Baillie and DeGennaro

(1990) finds that if one sets the conditional distribution of the stock return shock from normal to student's t, the positive direction of risk-return trade-off observed in French, Schwert and Stambaugh (1987) disappears. Recent studies propose better identification methodologies for detecting a positive relationship: focusing on sample size (e.g., Lundblad, 2007), measurement of expected return (e.g., Pastor et al., 2008), measurement of conditional variance (e.g., Ghysels et al., 2005; Jiang & Lee, 2014), regime-switching models (e.g., Ghysels et al., 2014; Nyberg, 2012; Salvador et al., 2014; Yu & Yuan, 2011) or the inclusion of a hedging factor (e.g., Guo & Whitelaw, 2006; Rossi & Timmermann, 2015; Scruggs, 1998).

3. Theory

This section discusses the application of standard finance theory to housing markets. First, I set up the theoretical prediction for empirical analyses. Second, I present clarifications for research in this area, considering the suitability of applying standard finance theory to housing assets.

3.1 Prediction

The empirical analysis of this research is based on Merton's (1973) ICAPM. Although one of the theoretical contributions in Merton (1973) is the existence of hedging phenomenon, the identity of the hedging factor has remained theoretically and empirically elusive in the literature. Merton (1980) contends that under certain conditions, the hedge component is negligible and the general intertemporal equilibrium mean-variance trade-off can still be a reasonable approximation. Therefore, related works (e.g., Bali, 2008; Frazier & Liu, 2016; Ghysels et al., 2005; Han, 2013; Lundblad, 2007; Nyberg, 2012; Pastor et al., 2008; Yu & Yuan, 2011) testing the intertemporal risk-return relationship often consider the reduced form of the ICAPM, assuming a constant investment opportunity in the future. The theoretical condition in Equations (1) and (3) can be further specified by the following empirical estimations:

$$R_{i,t} = c + (Risk\ Aversion)\sigma_{im,t} + \varepsilon_{i,t}, \tag{4}$$

$$R_{m,t} = c + (Risk\ Aversion)\sigma_{m,t}^2 + \varepsilon_{m,t}, \tag{5}$$

where $R_{i,t}$ and $R_{m,t}$ are the excess returns for a portfolio and the aggregate market, respectively; $\sigma_{im,t}$ is the conditional covariance between the portfolio risk premium and the aggregate market risk premium; $\sigma_{m,t}^2$ is the conditional variance for the aggregate market. If the theoretical prediction of a positive risk-return trade-off for risk-averse agents (i.e., the implied positive risk aversion) is true in Equations (4) and (5), it supports the validity of the practical application for the housing (or security) characteristic line:

$$R_{i,t} = \alpha_i + \beta_i\left[R_{m,t}\right] + \varepsilon_{i,t}. \tag{6}$$

Relaxing the single-period assumption about investors, Merton's (1973) ICAPM is built on continuous market equilibrium for multi-period consideration. Although the implication of a single-factor model is similar to Sharpe's (1964) and Linter's (1965) capital asset pricing model (CAPM), the major difference between the CAPM and the ICAPM is that the empirical investigation of Equations (4) and (5) has to be estimated through a time-series model (i.e., conditional analysis). In the housing literature, there is a general agreement (e.g., Beracha & Skiba, 2013; Cannon et al., 2006; Case et al., 2011) that investors, within a given time period, demand a higher return from housing assets that are riskier. However, there is no such consensus over whether this condition can be valid over time in Dolde and Tirtiroglu (1997), Miles (2008, 2011), Willcocks (2010), Morley and Thomas (2011), Han (2013), Karoglou et al. (2013), Lin and Fuerst (2014), Morley and Thomas (2016) and Zhou (2016). Against this background, the major contribution of this paper is to identify the source of this contradiction by concentrating on the modelling of conditional risk and, in doing so, to offer a simple and intuitive empirical analysis of intertemporal mean-variance efficiency in housing markets.

3.2 Clarification

There are two ways to examine the validity of standard finance theory: normative and positive analysis (Bodie, Kane & Marcus, 2014). Normative analysis explores the assumptions of the model, whereas positive analysis examines the predictions. Given the complexity of the real world, the oversimplified assumptions of the CAPM and ICAPM are acknowledged to be invalid. For example, Merton's (1973) ICAPM strictly assumes that there are a sufficient number of investors with comparable wealth levels so that each investor believes that he can buy and sell as much of an asset as he wants at the market price, with no transaction costs, taxes or problems with indivisibility of assets. Therefore, the tests of the model are mostly positive in the literature, examining the predictions or implications of the theory.

It has long been known that housing prices and interest rates have an inverse relationship. Therefore, systematic risk—the derived pricing factor in standard asset pricing theory—can play a significant role in housing assets, especially considering the effect of monetary policy. However, regarding the empirical investigation of housing assets, it should be noted that research in the equity literature examines the implications of standard finance theory and, therefore, typically works with large portfolios (e.g., 25 portfolios in Fama & French, 1993) rather than individual securities. Underlying this approach is the intuition that systematic risk might not always play an important role in some securities or small portfolios. Through diversification, systematic risk is essential in a large portfolio. Similarly, I also acknowledge that systematic risk might not always be the major pricing factor for individual housing assets or some regional markets. Nonetheless, for the major housing markets, systematic risk could reasonably be expected to play a major role in pricing. It is therefore surprising that major regional markets, such as Chicago, have been shown to demonstrate a negative risk-return relationship (e.g., Han, 2013; Karoglou et al., 2013).

Han (2013) regards the negative risk-return relationship in housing markets as the risk-return 'puzzle' that cannot be explained by the standard finance theory of Merton's (1973) ICAPM. In contrast, I suppose whether such a conclusion stems from the conditional risk measurement used in the housing literature, which can be considered unsatisfactory from both theoretical and statistical perspectives. First, standard finance theory implies that the pricing factor for regional markets should be systematic risk (e.g., Bayer et al., 2010). In contrast, the housing literature that investigates theoretical predictions focuses on the intertemporal relationship between market risk premium and 'total risk'. Second, sample sizes adopted in the housing literature (ranging from 48 to 264) are insufficient for a GARCH (1,1) model. Given the extensive estimation methodologies and intensive data mining, many of the discovered negative risk-return relationships could be considered 'significant' by accident. In sum, I remain agnostic as to whether the theorised relationship in the standard asset pricing model can be valid in housing markets. Without considering more reasonable risk measurements, I argue that most of the claimed findings in the housing literature against standard finance theory may be spurious.

4. Aggregate Analysis

This section examines the intertemporal relationship between the market risk premium and conditional variance based on the aggregate condition in the Merton's (1973) ICAPM. Unlike most of the housing literature, I consider an aggregate wealth portfolio of housing assets at the national, rather than regional, level.

4.1 Data

The data required for the empirical analysis includes the risk-free rate and housing returns in the US market. As is common, I use the short-term (one month) treasury bill rate from the Center for Research in Security Prices (CRSP) as a proxy for the risk-free rate. I use the national Case-Shiller home price index as my proxy for the performance of an aggregate market in the US. Consistent with related literature, I require my return measurement to be the first difference of the natural logarithm.[2] Further, since the univariate GARCH-M analysis is known to be sensitive to sample size, I use the historical US housing market record from January 1954 to December 2014 to maximise the time series content.[3]

Ideally, total return measurement, rather than capital gain measurement, should be employed. However, total return measurement in housing is infeasible (e.g., Han, 2013). Return measurement per capital gains has therefore been used in related work on housing markets. Under certain circumstances, such an approach to risk-return analysis

[2] The proposed return measurement is the capital gain measurement. Because data are unavailable for a total return measurement, this approach is widely used in the related housing literature (e.g., Beracha & Skiba, 2013; Cannon et al., 2006; Case et al., 2011; Cotter et al., 2015; Dolde & Tirtiroglu, 1997; Han, 2013; Karoglou et al., 2013; Lin & Fuerst, 2014; Miles, 2008, 2011; Morley & Thomas, 2011).

[3] Though the national Case-Shiller home price index starts at 1890, my sample period starts at 1954 for consistency. Before 1953, the data are yearly. During 1953, there are 11 monthly observations.

is acceptable in the equity literature (e.g., Lundblad, 2007), such as when historical total return data are not available for a long sample period. Table 3 presents the related summary statistics. The low returns in the housing market can be interpreted as the potential hedging benefits of real estate. Therefore, investors demand lower returns in this asset category.

Table 3. Summary statistics of monthly asset returns

	Mean (%)	SD (%)	Min (%)	Max (%)	Obs
Risk-free rate	0.37	0.25	0	1.35	731
Housing return	0.34	0.56	-2.28	2.03	731

Notes: The returns are calculated as: $r_t = ln(P_t/P_{t-1})$ where P_t is the price index at time *t*, from Jan 1954 to Dec 2014. I use the national Case-Shiller home price index for the construction of the housing return and the one month treasury bill rate for the risk-free rate.

4.2 Methodology

Econometric modelling investigating the relationship between market risk premium and conditional variance instinctively lends itself to GARCH-M methodology:

$$r_{m,t} - r_{f,t} = c + \alpha_M \sigma_{M,t}^2 + \varepsilon_t, \tag{7}$$

where $r_{m,t}$ is the market return, $r_{f,t}$ is the risk free rate and ε_t is the error term of mean zero with conditional variance $\sigma_{M,t}^2$. The assumption of risk-averse agents implies that $\alpha_M > 0$.

Without loss of generality for the variance equation, I allow different standardised variance component specifications for the evolution of conditional volatility: GARCH (Bollerslev, 1986), Threshold GARCH (Zakoian, 1994), Exponential GARCH (Nelson, 1991) and Component GARCH (Engle & Lee, 1999):

$$\text{GARCH(1,1)}: \sigma_{M,t}^2 = \delta_0 + \delta_1 \varepsilon_{t-1}^2 + \delta_2 \sigma_{M,t-1}^2, \tag{8}$$

$$\text{TGARCH(1,1)}: \sigma_{M,t}^2 = \delta_0 + \delta_1 \varepsilon_{t-1}^2 + \delta_2 \sigma_{M,t-1}^2 + \delta_3 D_{t-1} \varepsilon_{t-1}^2, \tag{9}$$

$$\text{EGARCH(1,1)}: \ln\left(\sigma_{M,t}^2\right) = \delta_0 + \delta_1 \left(\left|\frac{\varepsilon_{t-1}}{\sigma_{M,t-1}}\right|\right) + \delta_2 \ln\left(\sigma_{M,t-1}^2\right) + \ \delta_3 \left(\frac{\varepsilon_{t-1}}{\sigma_{M,t-1}}\right), \tag{10}$$

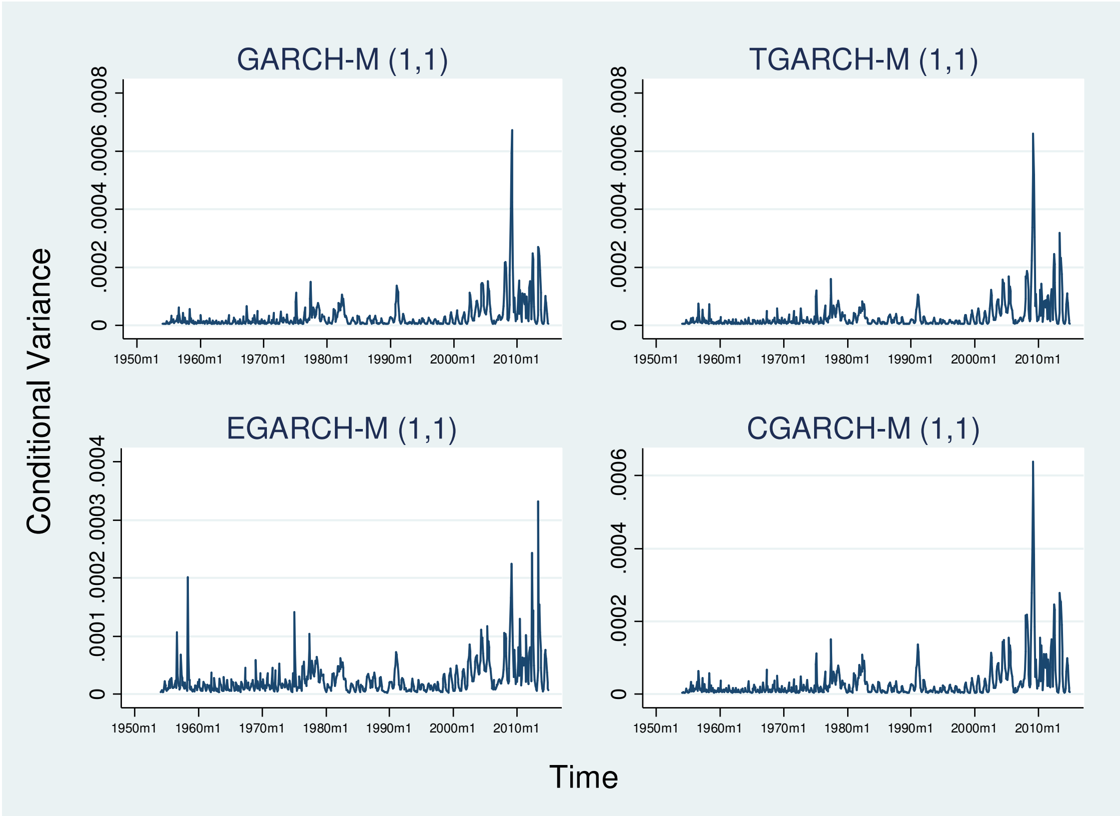
where D_t is a dummy taking the value of one when ε_t is negative and zero otherwise. The GARCH model suggests a symmetric response in conditional volatility after return innovation, while the δ_3 in the TARCH and EGARCH allows asymmetric shocks on conditional volatility (Engle & Ng, 1993). The variance equation of CGARCH takes a different form, allowing the volatility to be decomposed into permanent (q_t) and transitory ($\sigma_{M,t}^2 - q_t$):

$$\text{CGARCH(1,1)}: q_t = \delta_0 + \delta_1(q_t - \delta_0) + \delta_2\left(\varepsilon_{t-1}^2 - \sigma_{M,t-1}^2\right), \tag{11}$$

$$\sigma_{M,t}^2 - q_t = \delta_3(\varepsilon_{t-1}^2 - q_{t-1}) + \delta_4\left(\sigma_{M,t-1}^2 - q_{t-1}\right). \tag{12}$$

Using the return information, I obtain the predicted conditional variance using the GARCH-M models based on Equations (7) and (12). Figure 1 plots the predicted conditional variance from the baseline models with the normal error distribution. The conditional heteroskedasticity of variance is obvious in Figure 1. Several distinct periods of high market volatility (‘volatility clusters’) are visible. As expected, the housing market was more volatile during the recent Global Financial Crisis (GFC). Large volatility shocks seem to decay within a year across the entire sample.

Figure 1. Conditional variance generated by GARCH-Models



Note: This figure plots the conditional variance generated by GARCH-M models over the period Feb 1954 to Dec 2014.

4.3 Empirical Results

Because the results of GARCH-M models can be sensitive to volatility specification (Glosten, Jagannathan, & Runkle, 1993), I have adopted four common variants of GARCH-M models (GARCH-M, Threshold GARCH-M, Exponential GARCH-M and Component GARCH-M) for robustness check. Since Baillie and DeGennaro (1990) also show that results can be sensitive to the assumption about the distribution of error terms, I further consider three variants of error distribution for each model (normal, student's t and generalised). In total, I employ 12 versions of GARCH-M models with corresponding results reported in Table 4.

Table 4. GARCH-M results at the national level

	Model specifications	Error distribution	c	α_M (risk-return)	δ_0 ×10^{-6}	δ_1	δ_2	δ_3	δ_4	Log likelihood (adjusted R-squared)
(1)	*Panel A: GARCH(1,1)-M*	Normal	-0.001***	14.044***	3.91***	0.669***	0.211			2964.353
			(0.0002)	(5.424)	(0.562)	(0.096)	(0.045)			(-0.025)
(2)		Student's t	-0.001***	14.042***	3.90***	0.669***	0.211***			2964.353
			(0.0002)	(5.425)	(0.562)	(0.2105)	(0.045)			(-0.025)
(3)		Generalised	-0.0001***	15.905***	4.07***	0.597***	0.240***			2970.065
			(0.0002)	(5.427)	(0.508)	(0.063)	(0.040)			(-0.032)
(4)	*Panel B: TGARCH(1,1)-M*	Normal	-0.001***	24.764***	3.93***	0.821***	0.228***	-0.325**		2966.660
			(0.0002)	(5.982)	(0.571)	(0.157)	(0.049)	(0.160)		(-0.019)
(5)		Student's t	-0.001***	24.803***	3.94***	0.824***	0.227***	-0.328**		2966.618
			(0.0002)	(6.152)	(0.584)	(0.163)	(0.052)	(0.162)		(-0.018)
(6)		Generalised	-0.001***	22.177***	4.02***	0.686***	0.252***	-0.188*		2971.577
			(0.0002)	(5.839)	(0.500)	(0.092)	(0.043)	(0.111)		(-0.028)
(7)	*Panel C: EGARCH(1,1)-M*	Normal	-0.002***	39.844***	-3.70×10^{6}***	0.875***	0.727***	0.1497***		2958.495
			(0.0002)	(8.942)	(5.147)	(0.095)	(0.042)	(0.056)		(0.050)
(8)		Student's t	-0.001***	33.024***	-3.76×10^{6}***	0.899***	0.723***	0.1407**		2962.146
			(0.0002)	(8.047)	(0.502)	(0.104)	(0.043)	(0.057)		(0.046)
(9)		Generalised	-0.001***	12.963**	-3.75×10^{6}***	0.873***	0.723***	0.085*		2965.097
			(0.0002)	(5.603)	(0.440)	(0.080)	(0.038)	(0.045)		(0.023)
(10)	*Panel D: CGARCH(1,1) -M*	Normal	-0.0001***	12.975**	0.337	0.890***	0.659***	0.036	-0.488	2965.097
			(0.0002)	(5.600)	(0.234)	(0.081)	(0.105)	(0.039)	(0.445)	(-0.019)
(11)		Student's t	-0.0001***	12.964***	0.338	0.891***	0.660***	0.0364	-0.487	2965.074
			(0.0002)	(5.602)	(0.236)	(0.081)	(0.106)	(0.039)	(0.486)	(-0.019)
(12)		Generalised	-0.0009***	14.587***	0.254***	0.851***	0.581***	0.0476	-0.534*	2971.508
			(0.0002)	(5.575)	(8.77)	(0.058)	(0.072)	(0.0334)	(0.292)	(-0.024)

Notes: The table presents coefficient estimates based on Equations (7) to (12) for different versions of GARCH-M models. Numbers in parentheses denote the corresponding standard errors for the coefficients. ***, ** and * denote significance levels at 1%, 5% and 10%, respectively.

All of the GARCH-M models consistently show a significantly positive risk-return relationship based on the coefficient estimate of α_M, with 1% significance level in 10 of the 12 models and 5% significance level in the other two. This suggests that my result is robust in variance specifications and is not affected by the assumption about error terms. The consistency of the results allows me to draw the conclusion that there is a significantly positive risk-return trade-off in the national US market.

4.4 Implications

This study is the first to investigate the theoretically aggregate condition of Merton's (1973) ICAPM in the national housing market of the US. To better understand the implications of this approach and its success in examining the intertemporal risk-return relationship, I compare my results to previous literature on equity. The debate of ICAPM can be traced back to French et al. (1987) and Baillie and DeGennaro (1990), who find conflicting evidence regarding the sign of risk-return relationship in equity markets. Considering the empirical difficulty of documenting theoretical predictions via GARCH-M models, the subsequent literature devotes much effort to developing better identification methodologies (e.g., Frazier & Liu, 2016; Ghysels et al., 2005; Guo & Whitelaw, 2006; Jiang & Lee, 2014; Lundblad, 2007; Nyberg, 2012; Pastor et al., 2008; Rossi & Timmermann, 2015; Salvador et al., 2014; Scruggs, 1998; Yu & Yuan, 2011).

None of these studies, however, have considered the application of a different market proxy. Even though a focus on statistical estimation is important, the emphasis on the construction of an aggregate wealth proxy is no less paramount. The aggregate market proxy should ideally cover a wide range of capital assets, such as financial assets, real estate, government bonds, consumer durables and human capital (Fama & French, 2004). As highlighted in Roll's (1977) critique, a true market portfolio composed of all assets is most likely beyond reach. Hence, this study considers an alternative to the aggregate wealth portfolio by focusing on housing assets.

A problem common to all previous approaches is that the theoretically intertemporal positive risk-return relationship cannot robustly hold in equities across different types of GARCH-M models. In this study, I show that the empirical results of the intertemporal risk-return relationship in housing are surprisingly positive irrespective of volatility specifications in 12 versions of the standard GARCH-M models. The possible explanation for this success is that, unlike standard financial assets, housing is durable, serving as a long-term investment that provides a stream of services. The decision-making process for homebuyers is therefore known to be forward-looking (Paciorek, 2013). Such a factor indicates that the intertemporal risk-return relationship in housing can be more pronounced and thus more successfully detected by GARCH-M models.

5. Cross-Sectional Analysis

This section attempts to solve the risk-return puzzle at the regional level. I argue that the claimed risk-return puzzle could be an artefact of statistical modelling. I empirically explore this issue based on the cross-sectional condition in Merton's (1973) ICAPM.

5.1 Data

In this section, I attempt to solve the negative risk-return puzzle in a regional market. I choose Chicago, Cincinnati, Miami and Los Angeles as examples because Han (2013) and Karoglou et al. (2013) have found evidence for a significantly negative risk-return trade-off in these cities. The data for this empirical analysis include the risk-free rate and housing returns at the regional and national levels. I use the one month treasury bill rate from the CRSP as a proxy for the risk-free rate. Using the all-transaction housing price index data from the US Federal Housing Finance Agency (FHFA) over the period 1976:Q2 to 2015:Q4, I construct housing returns by using the first difference of the natural logarithm.[4] Table 5 displays the related summary statistics.

Table 5. Summary statistics of quarterly asset returns

	Mean (%)	SD (%)	Min (%)	Max (%)	Obs
Risk-free rate	1.18	0.88	-0.002	3.79	158
Housing return:					
Nation	1.08	1.25	-3.17	4.55	158
Chicago	1.05	1.97	-5.04	9.64	158
Cincinnati	0.86	1.30	-3.93	6.42	158
Miami	1.27	3.84	-19.36	18.94	158
Los Angeles	1.57	2.89	-8.14	10.21	158

Notes: The returns are calculated as: $r_t = ln(P_t/P_{t-1})$ where P_t is the price index at time *t* from 1976:Q2 to 2015:Q4. I use the FHFA housing price index for the construction of the housing returns and the one month treasury bill rate for the risk-free rate.

5.2 Construction of Conditional Covariance

Like Han (2013) and Karoglou et al. (2013), I first examine conditional risk modelling through the standard GARCH models. However, I highlight that the risk measurement for regional markets should be based on systematic risk, rather than the total risk generated by the univariate GARCH model. Therefore, following Bali (2008) and Bali and Engle (2010), I apply a bi-variate AR(1)-GARCH(1,1) model to construct the conditional covariance ($\sigma_{im,t}$):

$$R_{i,t} = \alpha_0^i + \alpha_1^i R_{i,t-1} + \varepsilon_{i,t}, \quad (13)$$

$$\sigma_{i,t}^2 = \gamma_0^i + \gamma_1^i \varepsilon_{i,t-1}^2 + \gamma_2^i \sigma_{i,t-1}^2, \quad (14)$$

[4] My sample starts at 1976:Q2 as the data are consistently available for the four cities.

$$R_{m,t} = \alpha_0^m + \alpha_1^m R_{m,t-1} + \varepsilon_{m,t}, \tag{15}$$

$$\sigma_{m,t}^2 = \gamma_0^m + \gamma_1^m \varepsilon_{m,t-1}^2 + \gamma_2^m \sigma_{m,t-1}^2, \tag{16}$$

$$\varepsilon_{i,t} * \varepsilon_{m,t} \equiv \sigma_{im,t} = \gamma_0^{im} + \gamma_1^{im} \varepsilon_{i,t-1}\varepsilon_{m,t-1} + \gamma_2^{im} \sigma_{im,t-1}, \tag{17}$$

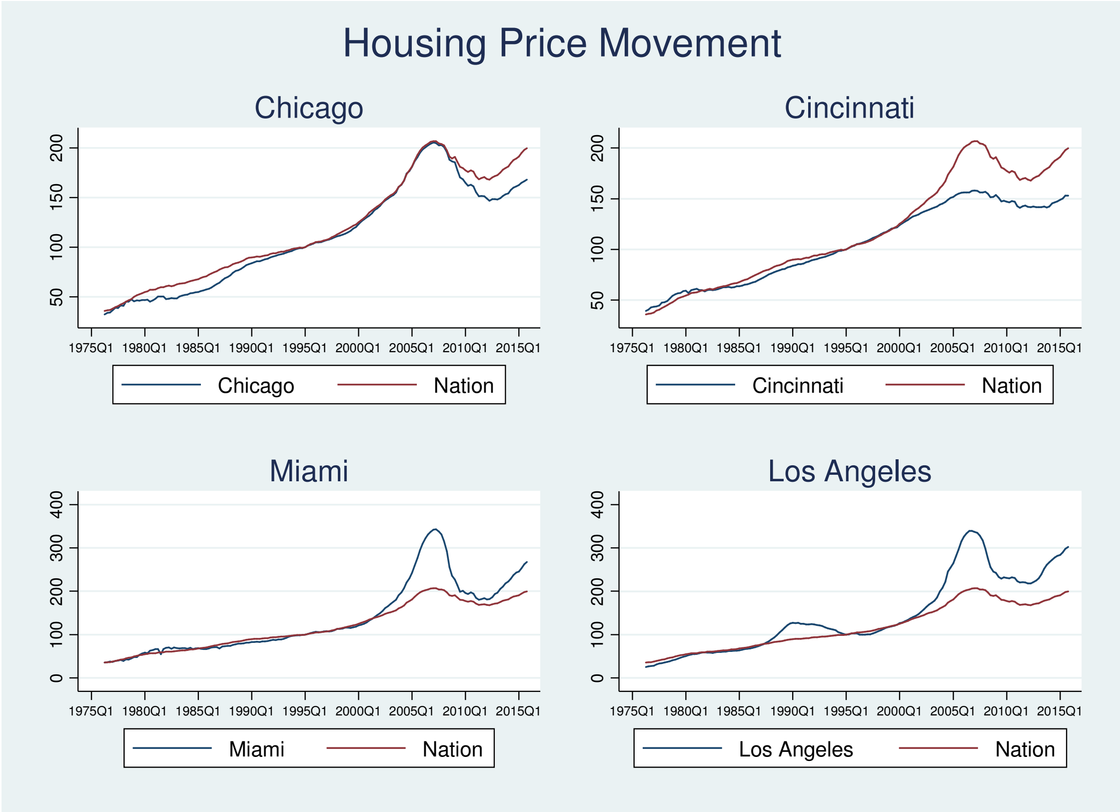
where $R_{i,t}$ and $R_{m,t}$ denote excess housing returns in the regional and national market at time (*t*), respectively; $\sigma_{i,t}^2$ and $\sigma_{m,t}^2$ represent the conditional variances for the regional and the national market, respectively; $\varepsilon_{i,t}$ and $\varepsilon_{m,t}$ are the error terms for the regional and the national market, respectively; $\sigma_{im,t}$ is the generated conditional covariance.[5]

5.3 First Stage Evidence

Using different types of GARCH-M models, the housing literature suggests a negative risk-return relationship in some regional housing markets, such as Chicago and Cincinnati in Han (2013) and Chicago, Miami and Los Angeles in Karoglou et al. (2013). To explain this finding, which contradicts the positive relationship postulated by standard asset pricing theory, Han (2013) argues that most variations in housing prices are local rather than national, whereas Karoglou et al. (2013) suggest that investors in the markets of a negative risk-return trade-off are less concerned with risk when making their investment decisions.

[5] The adopted programming of the covariance matrix is restricted to the general specification of BEKK (named after Baba, Engle, Kraft and Kroner) in Engle and Kroner (1995).

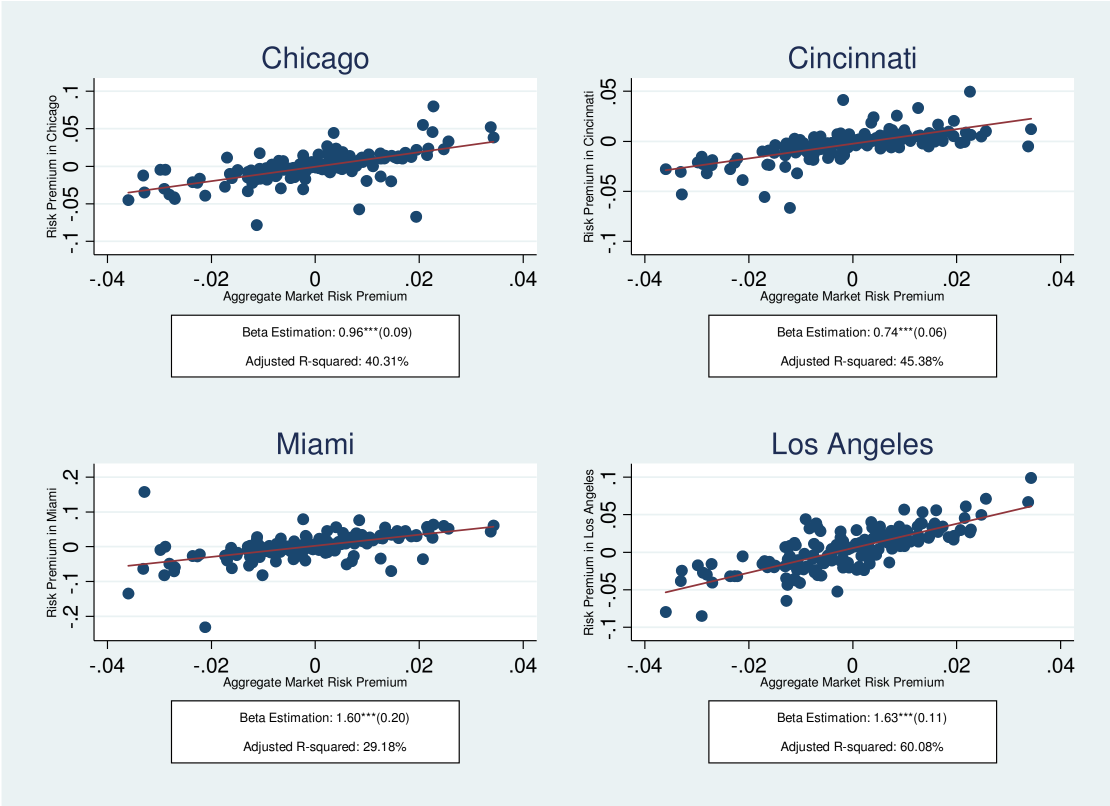
Figure 2. Housing price movement of puzzling markets



Notes: This graph displays the housing price index over the period 1976:Q2 to 2015:Q4. The data are collected from the US FHFA. To enable direct comparison of the two indices, I reconstruct the national price index using 1995:Q1 as the base quarter equivalent to 100.

Though these arguments offer potential explanations of the risk-return puzzle, the empirical evidence of housing price movements in Figure 2 and the housing characteristic line in Figure 3 present a different story. Based on the fundamental time series data analysis of regional markets against the national market in Figure 2, a pattern emerges that contradicts Han's (2013) argument. The trend of housing prices in Chicago, Cincinnati, Miami and Los Angeles is similar to the national level over the entire sample. In addition, analysis of the housing characteristic line in Figure 3 reveals that national market risk premium alone can account for 40.31%, 45.38%, 29.18% and 60.08% of the variations in the housing markets of Chicago, Cincinnati, Miami and Los Angeles, respectively. Contradicting Karoglou et al. (2013), the empirical result of significantly positive beta also suggests that investors in these cities are sensitive to systematic risk in the housing market.

Figure 3. Housing characteristic line



Notes: The figure depicts the results of the housing characteristic line. The model is specified as: $r_{i,t} - r_{f,t} = \alpha_i + \beta_i[\, r_{m,t} - r_{f,t}] + \varepsilon_{it}$ where α_i is the constant term, $r_{i,t}$ is the return of a regional housing market, $r_{m,t}$ is the return of the national market and $r_{f,t}$ is the risk free rate. The housing returns are calculated as: $r_t = ln(\, P_t/P_{t-1})$ where P_t is the all-transaction price index at time *t* from 1976:Q2 to 2015:Q4. The national and regional housing index is obtained from the US FHFA. The proxy for the risk free rate is the one month treasury bill rate from the CRSP. For the estimation, numbers in parentheses denote the corresponding standard errors for the coefficients. ***, ** and * denote significance levels at 1%, 5% and 10%, respectively.

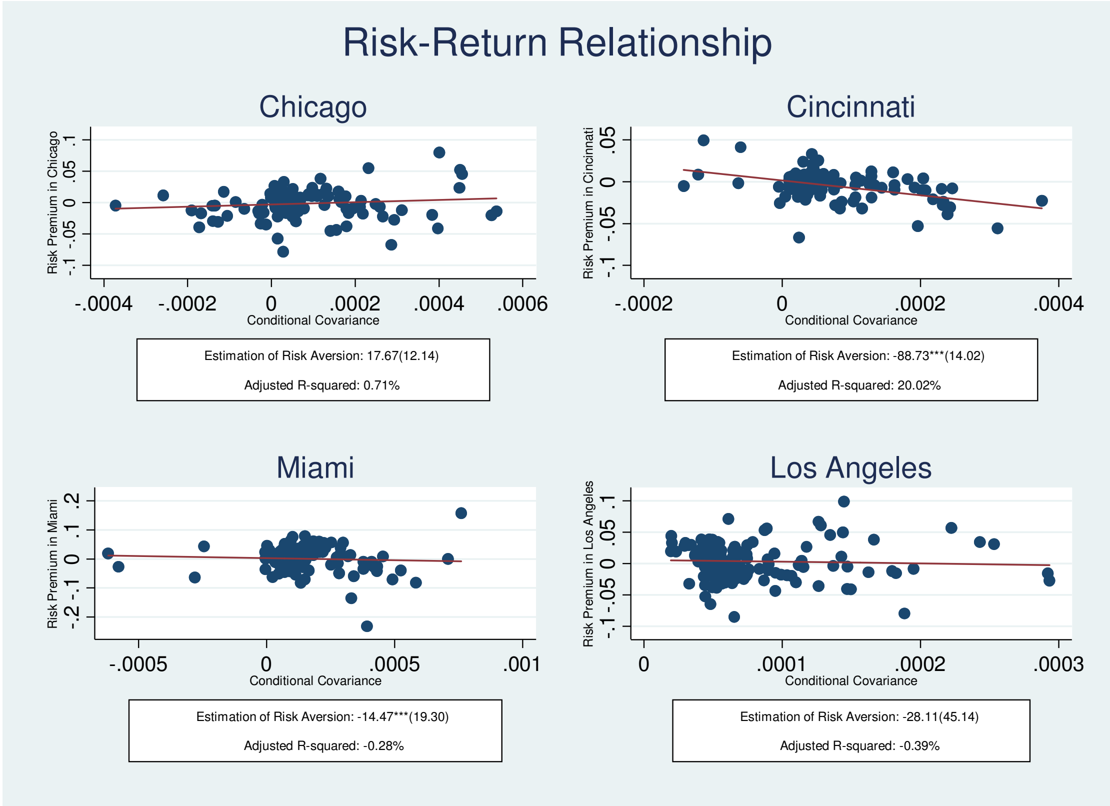
Given the illiquidity, transaction costs and information asymmetry present in property markets, it is plausible that variations in individual housing prices are local (i.e., idiosyncratic) rather than national (i.e., systematic). However, such a statement should not be generalised into the analysis at a sub-market level. The performance of the regional housing market can be interpreted as the aggregate performance of housing assets in one area. One can therefore view regional housing markets as comparable to large portfolios in the equity literature. From this perspective, systematic risk should play a significant role in a large, diversified portfolio composed of housing assets.

5.4 Second Stage Evidence

I next examine the risk-return relationship based on bi-variate GARCH models. My risk measurement here is similar to Han's (2013). However, I adjust the measure from the view of the cross-sectional condition in Merton's (1973) ICAPM for regional markets. Under this approach, as reported in Figure 4, I find an insignificant risk-return relationship in Chicago and Los Angeles, in contrast to the significantly negative relationship documented in the literature.

The recent works of Han (2013) and Loutskina and Strahan (2015) argue that financial analysis at the regional level should be conducted on samples taken before the GFC to avoid special circumstances. During the GFC, price movements in capital assets were unusual and unpredictable. By comparing the results of the full sample displayed in Figure 4 to the sub-sample taken before the GFC in Figure 5, I observe that, during the GFC, higher systematic risk was mostly associated with lower returns, which in turn weakened the positive risk-return trade-off for the overall sample. When excluding the crisis, I find a significantly positive risk-return trade-off in Chicago, with a significant improvement for model fitness of an adjusted R-squared value to 8.07%.

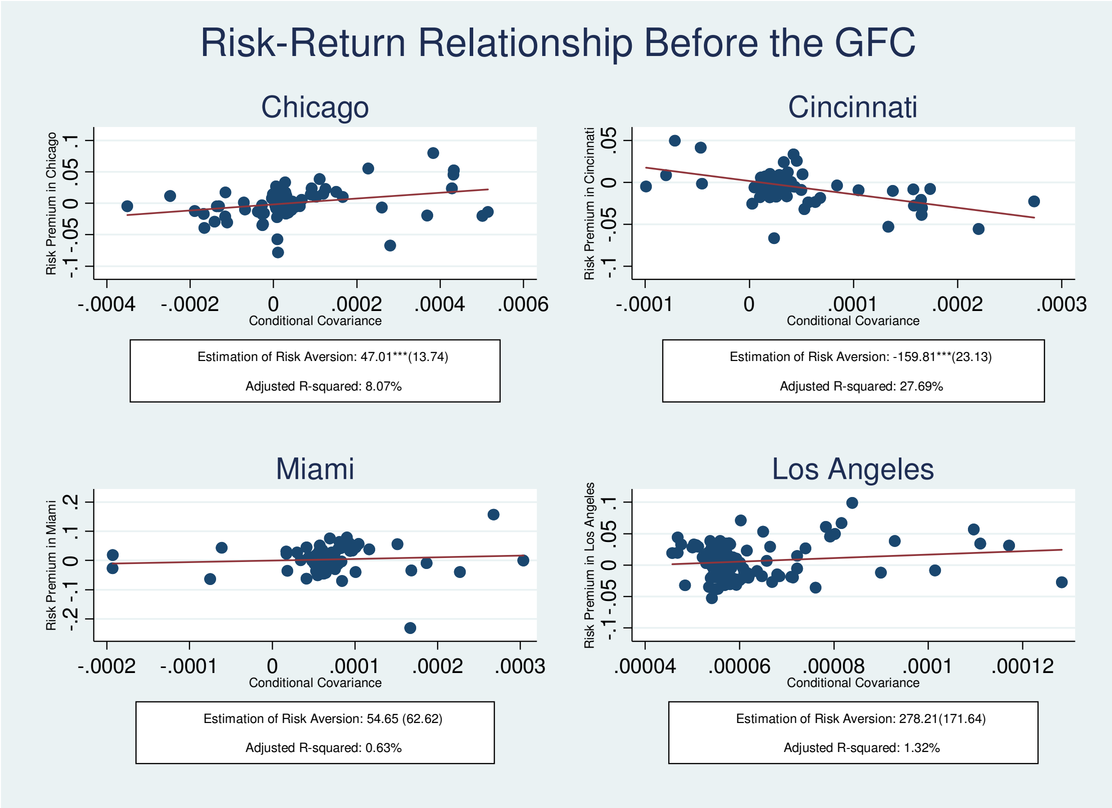
Figure 4. Risk-return relationship



Notes: The risk is generated by the bi-variate GARCH models based on Equations (13) to (17) from 1976:Q3 to 2015:Q4. With the predicted conditional covariance, I estimate Equation (4): $R_{i,t} = c + (Risk\ Aversion)\sigma_{im,t} + \varepsilon_{m,t}$. ***, ** and * denote significance levels at 1%, 5% and 10%, respectively.

Regarding the rest of the puzzling markets documented in the literature, I find an insignificant risk-return relationship in Miami and Los Angeles but a significantly negative relationship in Cincinnati. Even though the true risk-return relationship is positive, the observed negative relationship is not surprising since the literature recognises that GARCH models based on a short-time sample can produce unreliable estimates for conditional risk (e.g., Guo et al., 2013; Hill & Prokhorov, 2016; Hwang & Pereira, 2006; Iglesias & Phillips, 2011; Lundblad, 2007; Willcocks, 2009). My result highlights that a different estimation method with the application of a limited sample size can lead to an opposing conclusion.

Figure 5. Revised risk-return relationship before the GFC



Notes: The risk is generated by the bi-variate GARCH models based on Equations (13) to (17) from 1976:Q3 to 2007:Q2. With the predicted conditional covariance, I estimate Equation (4): $R_{i,t} = c + (Risk\ Aversion)\sigma_{im,t} + \varepsilon_{m,t}$. ***, ** and * denote significance levels at 1%, 5% and 10%, respectively.

5.5 Implications

Most empirical tests in the equity literature examine the implications of standard finance theory by focusing on large portfolios. If risk-averse investors can hold large diversified portfolios of assets, one of the research questions in Merton's (1973) ICAPM becomes whether there is a positive risk-return trade-off between systematic risk and the portfolio risk premium in an intertemporal framework. The negative risk-return trade-off in Chicago, as documented in Han (2013) and Karoglou et al. (2013), is problematic since it implies that most housing investors are risk-seeking. To resolve this puzzling phenomenon, I consider the setting of the cross-sectional condition in Merton (1973). I find a significantly positive risk-return trade-off in Chicago, supporting the assumption that investors are risk averse.

The fundamental time series illustration in Figure 2 clearly shows that when the national housing price increases, the regional housing price will move in the same direction. This suggests that if there is a higher level of systematic risk, there will be higher returns in regional markets. The documented negative risk-return relationship in some regional markets can be related to choices of estimation methods. Thus far, no consensus has been reached in the literature regarding the best method of modelling conditional risk, since econometricians cannot directly observe investors' information sets. This is especially problematic in the application to real estate markets, given the unestablished dataset, in contrast to financial markets. In this section, I do not argue that my proposed application of the bi-variate GARCH is the best method for modelling conditional covariance for regional markets. Yet, my analysis does imply that different modelling techniques can result in opposite conclusions. Thus, I urge that one should exercise caution when interpreting conflicting findings in the housing literature.

6. Further Cross-Sectional Analysis

Han (2013) documents an intertemporal negative risk-return trade-off across US regional markets through a panel data analysis. This section revisits this analysis, focusing on the cross-sectional condition of Merton's (1973) ICAPM.

6.1 Data

The data sources for the regional analysis are consistent with the previous section. I apply the one month treasury bill rate from the CRSP as a proxy for the risk-free rate. I use the all-transaction housing price index from the US FHFA to construct housing returns by calculating the first difference of the natural logarithm. Following the majority of related housing literature, the proposed sample period starts from 1985:Q1 to maximise the representation of US Metropolitan Statistical Areas (MSAs).[6] After dropping missing variables in some of the regional markets, the sample contains 221 MSAs.

My sample period ends in 2007:Q2 to avoid the possible confounding effects of the recent GFC. Excluding the GFC period from my sample allows a more direct comparison of my risk-return analysis, the conditional work of Han (2013) and the unconditional work in Cannon et al. (2006), Case et al. (2011) and Beracha and Skiba (2013). Finally, I repeat the bi-variate GARCH modelling procedure to construct conditional covariance measures for each of the regional markets. The resulting sample includes 19,448 market-quarter observations.[7]

6.2 Empirical Methodology and Results

To examine whether the intertemporal positive risk-return trade-off holds across sub-markets, I apply a panel data analysis to Equation (4) as follows:

$$R_{i,t} = c + (Risk\ Aversion)\sigma_{im,t} + \alpha_i + \beta_t + \varepsilon_{i,t}, \tag{18}$$

where $R_{i,t}$ is the excess return for each of the regional housing markets, $\sigma_{im,t}$ denotes the conditional covariance with the national market, α_i indicates regional fixed effects and β_t represents time fixed effects. Table 6 presents the results of this analysis, which reveals a significant and positive intertemporal risk-return relationship across regional markets. The coefficient estimates of conditional covariance (i.e., the implied risk aversion) remain

[6] Though some of the regional series are available from 1975, most of the series are well-established at a later stage. Hence, following Case et al. (2011) and Cotter et al. (2015), the proposed sample period starts at 1985:Q1.
[7] Regarding the 19,448 market-quarter observations (i.e., 88×221), 221 is the number of surveyed markets and 88 is the number of time series observations for each MSA after the bi-variate GARCH modelling.

significantly positive across Columns (1) to (4) at the 1% significance level. In contrast to Han (2013), my approach of highlighting the importance of systematic risk in the cross-sectional condition of standard finance theory reveals a significantly positive risk-return trade-off across regions and over time.

Table 6. Intertemporal risk-return relationship across MSAs

	(1)	(2)	(3)	(4)
Constant	-0.002***	-0.003***	-0.006***	-0.007***
	(0.0002)	(0.002)	(0.001)	(0.002)
Conditional covariance	75.850***	75.022***	38.508***	27.115***
	(3.133)	(3.441)	(2.991)	(0.337)
Regional fixed effects		✓		✓
Time fixed effects			✓	✓
Adjusted R-squared	2.92%	4.39%	25.39%	27.33%
Observations	19,448	19,448	19,448	19,448

Notes: This table presents the estimation of $R_{i,t} = c + (Risk\ Aversion)\sigma_{im,t} + \alpha_i + \beta_t + \varepsilon_{i,t}$ where $R_{i,t}$ is the excess return for each regional housing market, $\sigma_{im,t}$ denotes conditional covariance with the national market, α_i is a regional fixed effect and β_t is a time fixed effect. Conditional covariance is constructed through the bi-variate GARCH model over the period 1985:Q2 to 2007:Q2. Numbers in parentheses denote the corresponding standard errors for the coefficients. ***, ** and * denote significance levels at 1%, 5% and 10%, respectively.

6.3 Implications

Built upon the standard CAPM model with the assumption of a one-period investment horizon, a series of unconditional analyses in Beracha and Skiba (2013), Cannon et al. (2006) and Case et al. (2011) show that, within a given period, regional markets exhibiting higher systematic risk are associated with higher returns across MSAs. The assumption of fixed systematic risk is implicitly imposed in this strand of literature, and rarely have we considered a credible economic rationale for the assumption that the slope coefficients are time-invariant (Engle, 2015). In reality, the lives of investors span several periods and their investment decisions will vary over time based on their expectations (Bali, Engle, & Tang, 2016). Therefore, it is more intuitive to assume that the systematic risk of housing assets is dynamic rather than static. In this section, I consider a 'time-varying' risk-return trade-off in housing markets based on the cross-sectional condition of Merton's (1973) ICAPM. In contrast to Han (2013), my empirical investigation of the cross-sectional condition rather than the aggregate condition shows that the positive risk-return trade-off holds across sub-markets and over time.

7. Conclusion

Housing is an important asset class in the economy. However, the intertemporal nature of its risk-return relationship is not completely understood. While standard asset pricing theory typically predicts a positive relationship between risk and return, the empirical evidence reported in the extant literature shows conflicting signs of risk-return trade-off in housing markets. Identification of this empirical contradiction further leads Han (2013) to the conclusion that the risk-return puzzle cannot be explained by Merton's (1973) ICAPM. To address the puzzle in Han (2013), I consider the modelling differences in conditional risk between the aggregate and cross-sectional conditions in Merton's (1973) ICAPM for housing markets.

Based on the aggregate condition with the application of GARCH-M models, I find a significantly positive risk-return trade-off between market risk premium and conditional variance in the national market. While investigating the risk-return trade-off using GARCH models can be uninformative for short time series in regional markets, I consider whether the intertemporal risk-return trade-off can hold across regional markets in panel data models. I follow Han (2013) to first generate conditional risk measures for each regional market and pool the estimates into regression modelling. In contrast to Han (2013), I document a significantly positive risk-return trade-off between the market risk premium and conditional covariance based on the cross-sectional condition in standard finance theory.

I add to the empirical literature by explaining the conflicting evidence through the lens of aggregate and cross-sectional conditions in Merton's (1973) ICAPM. Unlike the majority of the housing literature (e.g., Dolde & Tirtiroglu, 1997; Han, 2013; Karoglou, Morley, & Thomas, 2013; Lin & Fuerst, 2014; Miles, 2008, 2011; Morley & Thomas, 2011, 2016; Willcocks, 2010; Zhou, 2016), my empirical results in the US housing markets confirm the prediction of standard finance theory that risk-averse agents require higher returns to reward higher risk in a multi-period setting. Altogether, my empirical findings regarding aggregate and cross-sectional analyses using Merton's (1973) ICAPM all yield an identical story—that housing markets can follow the classic mean-variance efficiency in an intertemporal framework. I therefore demonstrate that standard finance theory can resolve the argued risk-return puzzle in the extant housing literature.

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Paper 2. Mind the Gap: The Nature of Housing Markets

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Abstract

This article examines two theories about the nature of housing markets: first, much of the variation in housing price is local; second, time-invariant factors are important drivers of housing price dynamics. Through a variance decomposition analysis in the metropolitan statistical areas of the US from 1985 to 2015, I show that national factors, absorbed by year fixed effects, can account for 74.08% of variation in housing prices and 39.81% of variation in housing returns. The time-invariant factors proposed in the literature can be absorbed by regional fixed effects with low explanatory power of about 8.78% for housing prices and 0.24% for housing returns. Further analysis shows that the significance of time and regional fixed effects in panel data regression modelling is sensitive to the time-series dimension of the dataset. These results have implications for doctrinal assumptions made about housing markets in the literature.

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1. Introduction

Despite the importance of housing to the economy, surprisingly little is known about the nature of housing markets. Though standard asset pricing models suggest that asset prices are determined by national factors, conventional urban models indicate that real estate prices are determined by local factors. Recent literature by Hwang and Quigley (2006), Han (2013) and Glaeser, Gyourko, Morales and Nathanson (2014) advocates that local rather than national factors drive most variations in housing prices. Beracha, Gilbert, Kjorstad and Womack (2018) further conclude that the desirability of time-invariant local amenities is a significant channel to understand housing price dynamics. This paper aims to improve our understanding of the evolution in housing prices via exploring the validity of the arguments in Hwang and Quigley (2006), Han (2013), Glaeser et al. (2014) and Beracha et al. (2018).

The debate about the nature of housing markets can be traced back to a division in the theoretical literature between standard asset pricing and traditional urban models. The standard asset pricing models of Sharpe (1964), Linter (1965), Mossin (1966), Black (1972) and Ross (1976) argue that systematic risk is the major pricing factor for asset returns, suggesting that asset price changes are mainly national (which is common to assets everywhere). This theoretical prediction of a positive risk-return trade-off is supported by empirical evidence in the housing literature (e.g., Beracha & Skiba, 2013; Cannon, Miller, & Pandher, 2006; Case, Cotter, & Gabriel, 2011). In contrast, traditional urban models (Alonso, 1964; Roback, 1982; Rosen, 1979) contend that housing prices are determined by local income and amenities, suggesting that housing price changes are mainly regional (i.e., they affect all houses within one region, but nowhere else). Along with the long-held belief that the nature of housing markets is mainly local, the literature has focused on the link between housing markets and local factors, such as land share in

Davis and Palumbo (2008), land regulation in Gyourko, Saiz and Summers (2008), land supply elasticity in Saiz (2010) and local amenities in Albouy (2016).

While a small empirical literature on the evolution of housing markets exists, empirical evidence directly investigating long-held assumptions about the nature of housing markets is scarce. Specifically, is much of the variation in housing price driven by local, rather than national, factors, as assumed in Hwang and Quigley (2006), Han (2013) and Glaeser et al. (2014)? If so, how much closer do previously identified local determinants help us move towards a more complete understanding of housing price dynamics? Is it true that time-invariant local amenities are critical drivers for housing price dynamics, as highlighted in Beracha et al. (2018)? To answer these questions, I quantify the extent to which existing local determinants explain cross-sectional and time-series variation in housing markets by comparing their relative importance to that of national factors. In doing so, I am able to not only assess the progress of existing empirical work, but also characterise the major effects on housing markets.

This research begins by examining the long-standing implication in urban models that housing markets are mainly local in nature. Following Glaeser et al. (2014), I adopt time fixed effects as proxies for estimating the overall effect of national factors on housing price dynamics. Based on a balanced panel of 221 metropolitan statistical areas in the US from 1985 to 2015, I show that year (quarter) fixed effects account for 74.08% (74.42%) of variation in housing prices, which contrasts with the documented evidence showing year fixed effects contribute to 8% in Glaeser et al. (2014). Though Glaeser et al. (2014) conclude that the nature of housing markets is mainly local, the empirical evidence in this study presents a different story. Overall, the significance of time fixed effects in panel data regression for housing markets could be significantly higher than previously thought.

I next examine the conclusion of Beracha et al. (2018) that the time-invariant local amenities in Albouy (2016) can improve our understanding of housing price dynamics. To ascertain this issue, I use the standard income and population variables in Han (2013) as a baseline model for comparison. Through variance decomposition analysis across 196 metropolitan statistical areas in the US, I show that time-varying income growth and population growth can account for approximately 13.63% and 1.26% of variation, respectively, in housing return dynamics in balanced panel data from 1985 to 2015. In contrast, the commonly used time-invariant local factors in the housing literature, such as land regulation in Gyourko et al. (2008), land supply elasticity in Saiz (2010) and local amenities in Albouy (2016), each adds explanatory power of no more than 1.54% to the baseline model. This result contradicts the evidence of Beracha et al. (2018) that the time-invariant amenity of a metropolitan area is significant for understanding house price dynamics.

To shed light on why the empirical evidence in this paper contradicts Glaeser et al. (2014) and Beracha et al. (2018), I explore the significance of time and regional fixed effects in panel data regression modelling with the application of different time series dimensions to the dataset. My further analysis shows that the documented evidence for relatively lower explanatory power of time fixed effects (i.e., a proxy for time-varying national factors) yet higher explanatory power of regional fixed effects (i.e., a proxy for time-invariant regional factors) can be obtained by reducing the time-series dimension of data samples. Estimating panel data models with fixed effects typically requires at least two or more time periods. In an extreme case of a one period sample, time fixed effects are not applicable and thus one can consider that its contribution to the variation of housing price is zero (i.e., 0% R-squared). Since there is also no time series variation in a one period sample, regional fixed effects identify that all local factors are time-invariant and thus lead to 100% R-squared. In contrast, the empirical evidence of a full sample from 1985:Q2 to 2015:Q4 shows that quarterly fixed effects alone can account for 74.10% of variation in housing prices and 31.70% in housing returns, while regional fixed effects alone can account for 8.78% of variation in housing prices and 0.24% in housing returns. Thus, my work cautions that the documented importance of time-varying national factors and time-invariant local factors can be sensitive to the selection of the time series dimension in modelling.

This work is mostly relevant to Fairchild, Ma and Wu (2015). Based on the dynamic asset pricing model of Campbell and Shiller (1988), Fairchild et al. (2015) investigate the relative importance of common national factors and local factors in housing price 'volatility' in each of the 23 US cities, finding that a large fraction of housing market volatility is local for most of these cities. This work expands on their research by focusing on decomposing the growth of housing wealth (i.e., housing 'returns') through the analysis of covariance in a more comprehensive panel data regression with fixed effects for 196 metropolitan statistical areas. This approach also advances understanding of the 'overall' effect of national factors on all regional housing markets.

Altogether, the effect of national economic conditions on local housing markets remains important in the US over the sample period of 31 years. In contrast to Hwang and Quigley (2006), Han (2013) and Glaeser et al. (2014), this study cannot conclude, in favour of the standard urban models, that housing price changes should be mainly local rather than national. My empirical evidence echoes Leung's (2004) compendium that conventional housing and urban economics research should incorporate interactions between housing markets and the macro-economy into theoretical and empirical modelling. The structure of the paper is as follows: Section 2 relates my work to two theories of housing price dynamics, Section 3 investigates the nature of housing markets, Section 4 explores the significance of time-invariant local amenities in housing markets, Section 5 presents further analysis and Section 6 concludes.

2. Literature Review

This section discusses two strands of literature that explain housing price changes. Next, I synthesise the two schools of thoughts and propose the argument of this research.

2.1 Standard Asset Pricing Theory

The canonical Capital Asset Pricing Model (CAPM) of Sharpe (1964), Linter (1965), Mossin (1966) and Black (1972) and the Arbitrage Pricing Theory (APT) of Ross (1976) highlight that asset returns are determined by exposure to systematic risk (i.e., the risk inherent in the entire market). Though the empirical tests of standard asset pricing theory have long centred on equity assets, Roll's (1977) critique suggests the application of real estate should be included. Therefore, as is usual for financial assets, the risk premium of housing assets should be their exposure to systematic risk (Bayer, Ellickson, & Ellickson 2010). This is empirically supported by Cannon et al. (2006), Case et al. (2011) and Beracha and Skiba (2013).

Systematic risk is particularly important in the housing market, especially considering the role of monetary policy in housing (e.g., Aoki, Proudman, & Vlieghe, 2004; Goodhart & Hofmann, 2008; Rahal, 2016; Rubio, 2014). To illustrate, it is known that there is an inverse relationship between housing and interest rates based on conventional monetary policy. Under the influence of low interest rates, housing in the US started to jump to a 25 year high by the end of 2003 and remained high until a sharp decline began in early 2006 (Taylor, 2007). Evidence of unconventional monetary policy, in the form of innovations of the monetary base, also shows the positive and persistent response of house prices to policy shocks of increasing innovation (Rahal, 2016).

2.2 Standard Urban Model

Though standard asset pricing theory implies no role for idiosyncratic risk in the asset pricing of housing markets (i.e., the local risk within a region), local factors can be important in regional housing markets. The standard urban models of Alonso (1964), Rosen (1979) and Roback (1982) propose that housing prices reflect a spatial equilibrium, in which prices are determined by local wages and amenities. Specifically, urban models typically formulate the housing decision of where to live as a discrete choice about a bundle of housing and neighbourhood attributes.

The local 'fundamentals' driving the change in house prices have long been emphasised in the housing literature. Hwang and Quigley (2006) and Beracha et al. (2018) provide a complete review emphasising and identifying local economic fundamentals in housing markets, confirming the significance of regional economic conditions to local markets. To further advance our knowledge of local dynamics, several related indices are proposed and published, such as the land share value in Davis and Palumbo (2008), the land regulation index in Gyourko et al. (2008), land supply elasticity in Saiz (2010) and the quality of life, trade-productivity and total amenity indices in Albouy (2016).

2.3 Issue

While both standard asset pricing and urban models have strengths and weaknesses in explaining housing price dynamics, Hwang and Quigley (2006), Han (2013) and Glaeser et al. (2014) have reached the consensus that the variation of housing prices is mainly local rather than national. A recent study by Bercha et al. (2016) further argues that the local time-invariant amenities of Albouy (2016) are critical for understanding housing price dynamics, especially when previous research utilising macroeconomic variables does not adequately explain the variations in housing price dynamics across Metropolitan Statistical Areas (MSAs). All of this suggests a less prominent role is played by national factors in housing markets, supporting standard urban over standard asset pricing models.

2.4 Proposed Argument

Fundamentally, housing prices are all affected by common aggregate factors based on standard asset pricing models. Yet, local supply and demand can drive heterogeneity in housing price changes across cities, according to conventional urban models. In contrast to Hwang and Quigley (2006), Han (2013) and Glaeser et al. (2014), who conform to the view that variation in housing price is mainly local rather than national, I argue that the relationship between standard asset pricing and urban models is complementary rather than competing in understanding housing dynamics.

This paper seeks to better understand the nature of the shocks driving fluctuations in regional housing markets and further assesses the validity of proposed local factors through a balanced panel data model with fixed effects across 196 MSAs in the US from 1985 to 2015. As highlighted in Fairchild et al. (2015), quantitatively separating the effect of national factors from that of local factors in the housing markets is essential, since policymakers, for example, are interested in whether monetary policy is responsible for local housing bubbles by keeping the short-term interest rate low for a long period of time. By further comparing the relative importance of proposed local factors, this paper also contributes to the literature by providing a deeper understanding of how local factors drive housing price dynamics.

3. Empirics: Nature of Housing Markets

The key method for investigating the nature of housing markets is the analysis of covariance (ANCOVA). By quantifying the explanatory power of time fixed effects for the total variation in housing, I investigate whether the nature of housing markets is mainly local or national.

3.1 Research Methodology and Data Description

My empirical work is built upon a parametric ANCOVA, which allows me to decompose the variation in housing price into different factors. I do so by estimating the following panel data regression and then quantifying the explanatory power of the time fixed effects:

$$Housing\ Market_{i,t} = c + \beta_t + \varepsilon_{i,t}, \quad (1)$$

where $Housing\ Market_{i,t}$ is one of three time-series housing market variables: price, return (i.e., the first difference of the natural logarithm) and risk premium (i.e., asset return minus the risk-free rate) for the dependent variable and β_t is a time fixed effect. To directly investigate the nature of housing markets, I only consider the time fixed effect as the major independent variable. The unobserved effect of common macro factors at the national level, such as changes in interest rates and in tax codes, are absorbed by time fixed effects in a panel analysis. Therefore, the housing literature of Han (2013) and Glaeser et al. (2014) views time fixed effects as a proxy for national factors. This approach enables me to directly decompose the relative importance of national and regional factors in housing markets. To illustrate, if time fixed effects account for 10% of the variation in housing price dynamics, this indicates that national factors account for 10% of the variation.

To estimate the panel data regression, I obtain the all-transaction housing price index from the US Federal Housing Finance Agency (FHFA). To maximise representation of US MSAs, my proposed sample period spans from 1985:Q1 to 2015:Q4.[1] In total, there are 221 MSAs housing price indices for a balanced panel data regression. In my analysis, I consider three types of series in housing markets: price, return and risk premium. The risk premium series is of particular relevance since, fundamentally, standard asset pricing theory assumes that investors can borrow and lend at a risk-free rate. I use the one month treasury bill rate from the Center for Research in Security Prices (CRSP) as a proxy for the risk-free rate.

3.2 Results

Based on the results of the price series, I find that quarterly time fixed effects, on average, can account for 74.42% of the variation across regional housing markets. The result is at odds with Glaeser et al. (2014), who find that time fixed effects can only

[1] The beginning of the sample period follows Case et al. (2011) and Cotter, Gabriel and Roll (2015). Though FHFA provides a regional housing price index from 1975, most of the series are well-established at a later stage. I also drop the regional housing index that contains missing variables after 1985.

contribute to 8% of variation in housing prices. When considering the housing return series, the adjusted R-squared decreases from 74.42% to 31.70%. The result is not surprising, since the first difference method used to compute returns can make data more stationary by removing common trends, such as seasonality. Similar to the results for variation in housing returns, time fixed effects can explain 32.43% of variation in housing risk premiums.

The first difference of a time series (i.e., the series of changes from one period to the next) is usually considered in housing modelling. Using non-stationary time series data in modelling can produce unreliable and spurious results, leading to poor understanding and forecasting between dependent and independent variables. To illustrate, the results obtained by using non-stationary time series can be spurious since the relationship between dependent and independent variables may be driven by other common macro factors. To remove the unknown common trend, one common practice in time series analysis is to first-difference the housing price series. Therefore, it is reasonable to find that the effect of national factors with significantly high adjusted R-squared at the housing price series of 74.42% reduces to 31.70% of the housing return series.

Table 1. Impact of quarter fixed effects on housing

	Price	Return	Risk premium
Independent variable (quarter fixed effects)			
Adjusted R^2	74.42%	31.70%	32.43%
Observations	27,404	27,183	27,183
(time series observations×markets)	(124×221)	(123×221)	(123×221)

Notes: Table 1 presents the result of adjusted R-squared based on Equation (1): $Housing\ Market_{i,t} = c + \beta_t + \varepsilon_{i,t}$ with the application of quarterly data from 1985:Q1 to 2015:Q4.

I next consider whether the different conclusions drawn in the literature about the nature of local housing markets could result from sampling differences in time-series frequency. Consistent with Glaeser et al. (2014), I model my panel data regression based on the annual series. Following Han (2013), I use annual values imputed in the third quarter to represent the expected values for that year. My result of 74.08% of variation in housing prices explained by year fixed effects is, again, in sharp contrast to Glaeser et al. (2014), who find that only 8% of the variation in housing prices can be attributed to year fixed effects.[2] Combining the results of Tables 1 and 2, whether the nature of housing markets is mainly local or national depends on whether our dependent variable in panel data models is the housing price, return or risk premium. However, despite the difference in the adjusted R-squared explained by national factors across three types of series in housing markets, national factors play an important role in regional markets: at least 31.70% based on the modelling of quarterly series and 39.81% based on that of annual series.

Table 2. Impact of year fixed effects on housing

	Price	Return	Risk premium
Independent variable (year fixed effects)			
Adjusted R^2	74.08%	39.81%	39.89%
Observations	6,851	6,630	6,630
(time series observations×markets)	(31×221)	(30×221)	(30×221)

Notes: Table 2 presents the result of adjusted R-squared based on Equation (1): $Housing\ Market_{i,t} = c + \beta_t + \varepsilon_{i,t}$ with the application of annual data from 1985 to 2015.

3.3 Implications

The housing literature typically states that the nature of housing markets is mainly local rather than national (e.g., Glaeser et al., 2014; Han, 2013; Hwang & Quigley, 2006). Therefore, outcomes in housing markets reflect regional economic conditions. In this section, I show that this presumption might not always hold true, since the empirical

[2] Regarding the documented 8% of variation in annual housing prices attributed to national factors in Glaeser et al. (2014), the data sample is not clearly specified. Therefore, I follow Han (2013) and use annual values in the third quarter to represent the expected values for that year.

evidence about the nature of housing markets can vary, depending on the variables of interest. Based on the evidence of housing price series from 1985:Q1 to 2015:Q4, I find that national factors, characterised by quarter fixed effects, can account for approximately 74.42% of variation in the MSAs of the US. In this case, much of the variation in housing prices is national, which challenges the view of Han (2013, p. 878), '*Unlike other financial assets, much of the variation in house prices is local, not national*'.

Based on the evidence of the return and risk premium series, I find that time fixed effects can explain approximately 32% to 40% of variation in regional housing markets, depending on the frequency of the time series data. In this case, though much of the variation in housing markets is local, national factors remain significantly important for understanding housing price dynamics. Therefore, the argument of Hwang and Quigley (2006), Han (2013) and Glaeser et al. (2014)—that the nature of housing markets is mainly local—should more reasonably be viewed as a presumption rather than a fact.

4. Empirics: Importance of Local Factors

After disclosing the nature of housing markets, this section investigates the relative importance of local factors for housing price dynamics.

4.1 Data

The essential elements in the standard urban models of Alonso (1964), Rosen (1979) and Roback (1982) are local amenities and income. Following Beracha et al. (2018), I first collect the main independent variables of quality of life (i.e., amenities for households), trade productivity (i.e., amenities for firms) and total amenity indices (i.e., a combination of the amenities for both households and firms) in Albouy (2016) as controlling factors for local amenities. I then match these indices with the housing price index from FHFA. I remove markets with missing variables and focus on the sample from 1985 to 2015 to maximise the representation of US MSAs. Following Han (2013), I use annual values in the third quarter to represent the expected values for that year. Finally, I eliminate the observations for the markets in which the values of standard fundamentals for housing in Han (2013) (e.g., income and population, as discussed below) are missing. In total, I have a main sample of 196 regional markets with three groups of explanatory variables, whose summary statistics are reported in Table 3.

For the first group of census controlled variables, I obtain annual statistics from 1985 to 2015 of income and population from the Bureau of Economic Analysis in the US Department of Commerce to control the demand side of housing. I measure the income and population growth by taking the first difference of the natural logarithm. The average population growth per year is around 1% in MSAs, whereas the average income growth per year is around 5%. A view of either an ocean or a lake can potentially affect housing values (Landry & Hindsley, 2011; Wyman, Hutchison, & Tiwari, 2014). Thus, I follow Beracha et al. (2018) and create a dummy indicator that assigns a value of 1 if a city is situated within 70 miles of an ocean or within 40 miles of the Great Lakes (using Google Maps). Summary statistics indicate that around 39% of the 196 regional markets in my sample are close to the ocean and the Great Lakes.

For the second group of variables, I consider three measurements of amenities from Albouy (2016). The main hypothesis in Beracha et al. (2018) is that the desirability of an area from the perspective of homeowners (measured by quality of life) and firms (measured by trade productivity) in Albouy (2016) is an important channel through which land values affect house price dynamics. Combining the value of trade productivity with the value of quality of life from Albouy (2008), Albouy (2016) measures each city's total amenity value by incorporating non-residential land, local labour costs and federal tax externalities. Higher value of the indices indicates relatively higher quality of life, trade-productivity and total amenities. For instance, preliminary analysis in my sample shows that San Francisco is the market with the highest total amenity value (0.32), whereas Santa Barbara has the second highest value (0.26).

Table 3. Summary statistics

Variables	Mean	SD	Min	Max	Markets	Obs
Housing returns	0.032	0.060	-0.573	0.364	196	5880
Census data:						
Coastal/lake	0.39	0.49	0	1	196	5880
Income growth	0.05	0.03	-0.36	0.42	196	5880
Population growth	0.01	0.01	-0.29	0.09	196	5880
Albouy (2016):						
Quality	-0.0004	0.05	-0.10	0.21	196	5880
Productivity	-0.037	0.09	-0.22	0.29	196	5880
Total amenity	-0.02	0.09	-0.17	0.32	196	5880
Other published determinants:						
Land supply elasticity in Saiz (2010)	1.95	0.98	0.60	5.45	89	2670
Land share in Davis and Palumbo (2008)	0.37	0.21	0.05	0.89	42	1260
Land regulation in Gyourko et al. (2008)	-0.07	0.61	-1.13	2.32	196	5880

Note: Variables from 1986 to 2015 for population growth, income growth and land share are time-varying, whereas the other variables are time-invariant.

For published local determinants in the housing literature, I consider land share in Davis and Palumbo (2008), land regulation in Gyourko et al. (2008) at the state level and land supply elasticity in Saiz (2010). Using survey data, Gyourko et al. (2008) measure stringent building restrictions. A higher value for land regulation indicates more stringent building restrictions. Using geographic information system (GIS) techniques, Saiz (2010) estimates the slope and elevation of every 90 square metre parcel of land within 50 kilometres of the centroid of the area and computes the fraction of the undevelopable land in each MSA. A relative lower value for supply elasticity indices indicates less undeveloped land. In contrast to major published indices in the literature, land share in Davis and Palumbo (2008) is a time-varying and straightforward measure attained by computing the ratio of land value to total property value. Finally, I match these available published indices to my major sample of 196 markets. The annual summary statistics from 1986 to 2015 are reported in Table 3.

4.2 Model Specification

Consistent with the empirical analysis in Section 3, my work is built on a parametric framework, specifically, the ANCOVA to decompose the variation in housing returns into different factors. I first estimate the expanded version of Equation (1) and then quantify the explanatory significance of each independent variable:

$$Housing\ Return_{i,t} = c + \gamma X_{i,t} + \alpha_i + \beta_t + \varepsilon_{i,t}, \tag{2}$$

where I focus on the housing return as the dependent variable, $X_{i,t}$ is a group of local factors (detailed in Section 4.1), α_i is a regional fixed effect and β_t is a time fixed effect. The unobserved effect of common macro factors at the national level can be absorbed by time fixed effects in a panel analysis, while the regional fixed effects can control the permanent difference across cities, such as geographical constraints and local amenities (Han, 2013).

4.3 Empirical Results

Table 4 presents the empirical relationship between economic fundamentals and housing price dynamics. Column (1) in Table 4 is the baseline model with census data, in which I find, consistent with Han (2013) and Favara and Imbs (2015), that higher population growth and higher income growth are associated with higher housing returns. Han (2013) argues that income and population variables are major economic and demographic factors to control for the demand side of housing. Expanding this baseline model in Column (1) by including quality of life, trade-productivity and total amenities from Column (2) to Column (4), I find that there is a statistically positive relationship between housing returns and the three sets of amenity variables in Albouy (2016), reconfirming the empirical evidence in Beracha et al. (2018) that markets with higher amenities are associated with higher returns.

It should also be noted that the three measures of amenities carry similar information. Therefore, to avoid multi-collinearity, I follow Beracha et al. (2018) and only consider one amenity variable at a time in my modelling. If one adds the three sets of variables together in hedonic modelling, then none of the proposed amenity indices can explain the housing price dynamics. Indeed, most of the proposed indices in the housing literature are fundamentally computed from one to another and thus can carry similar information. For instance, to compute the amenity values, Albouy (2016) incorporates the land regulation of Gyourko et al. (2008) to control for unobserved housing productivity differences. Therefore, I expand the baseline model in Column (1) with only one additional explanatory variable at a time for the remainder of the analysis.

According to the homevoter hypothesis in Fischel (2001), homeowners have stronger incentives to protect their housing investments if land values were initially higher. Saiz (2010) further argues that local land use controls can be mainly understood as instruments for local homeowners to maximise land prices. Consistent with this prediction, Column (5) indicates that markets with higher land regulation are associated with higher housing returns. As highlighted in Gyourko et al. (2008), housing markets with stricter regulation are often situated in coastal areas. In Column (6), I further find that coastal markets tend to have higher returns. Although the literature of Landry and

Hindsley (2011) and Wyman et al. (2014) based on the property level analysis suggests a premium for waterfront views, Case et al. (2011) show that coastal markets have relatively higher systematic risk that could also justify higher returns.

Housing markets with lower land supply elasticity concentrates on coastal areas (Saiz, 2010). The theoretical model in Paciorek (2013) shows that regulations can lower the elasticity of the housing supply. Therefore, it is expected that markets with relatively higher elasticity will have lower returns. Though this negative relationship is not significantly captured in Beracha et al. (2018), it is statistically significant in Column (7). Furthermore, properties are typically capitalised into the value of land but not the value of the physical structure on the parcel of land, suggesting that changes in overall property value will depend critically on how much of its value is represented by land value (Bostic, Longhofer, & Redfearn, 2007; Davis & Heathcote, 2007).[3] Consistent with this theoretical prediction, I find that markets with higher values of land share are associated with higher returns.

4.4 Variance Decomposition Analysis

I now focus on the relative importance of fundamentals in housing markets by applying a parametric framework, specifically ANCOVA, by decomposing the variation in housing returns attributable to different factors. Table 5 displays the results of variance decompositions for several specifications. Each column in Table 5 corresponds to a related model specification for housing returns in Table 4. I divided the partial sum of squares for each effect by the aggregate partial sum of squares across all effects in the

[3] Though Column (8) in Table 4 presents an insignificant coefficient for population growth, this issue is resolved after controlling for time and regional fixed effects, which result in a significant improvement in model fitness.

model. This offers a normalisation that forces columns to sum to one. This value in the table represents the fraction of the model attributable to a specific effect.

The baseline model in Column (1) of Table 4 shows that the combined effect of income and population growth accounts for 14.89% of variation in housing returns. Further decomposition analysis in Table 5 shows that income growth exhibits explanatory power of 13.63% (which is calculated as 91.57% × 14.89%), while population growth does so to the tune of 1.26% (which is calculated as 8.43% × 14.89%). The expanded models, Column (2) to Column (7), which incorporate time-invariant local factors, show little additional explanatory power—at most 1.54% compared to the baseline model. In contrast, the expanded model in Column (8), which incorporates a time-varying land share index, significantly improves explanatory power by 7.54%.

This result contradicts the viewpoint of Beracha et al. (2018) that time-invariant factors, such as local amenities in Albouy (2016), significantly drive variation in house price 'dynamics'. Instead, I show that the time-varying land share index in Davis and Palumbo (2008) is another alternative to understanding housing price dynamics. Indeed, the effect of time-invariant local factors can be captured by regional fixed effects in panel data regression modelling. I will investigate this topic in greater detail in Section 5. In Columns (9) and (10) of Table 5, I find that time fixed effects, or a proxy for national factors, is the main determinant of housing price dynamics among all proposed factors.

4.5 Implications

Most of the proposed indices, such as land regulation in Gyourko et al. (2008), land supply elasticity in Saiz (2010) and amenity in Albouy (2016), are time-invariant and provide limited additional explanatory power compared to the baseline model.

Among the proposed indices, the land share index in Davis and Palumbo (2008) is particularly worthy of attention. This index is time-varying and also provides a relatively higher explanatory power for the local housing price dynamics (up to approximately 7.54% compared to the baseline model).

Income and population growth in the baseline model can account for approximately 13.63% and 1.26% of variation in housing return dynamics, respectively. When reviewing the relative importance of all variables in Columns (9) and (10), national factors, captured by time fixed effects, remain the most important, accounting for about 70% of the explained variation. Therefore, to better understand housing price ‘dynamics’, it is essential for Beracha et al. (2018) to control for time fixed effects.[4] When a time-series dimension is expanded to the dataset, my result contradicts the argument and empirical evidence of Beracha et al. (2018) that time-invariant amenity in Albouy (2016) can significantly improve our understanding of housing price ‘dynamics’ by showing that three types of amenities in Albouy (2016) have little explanatory power in the cross-sectional and time-series variation of housing returns.

[4] In a standard panel data regression with regional fixed effects, the model can remove any invariable difference across regional markets, such as geographical constraints and local amenities (Han, 2013; Loutskina & Strahan, 2015). This econometric fact also directly casts doubt on Beracha et al.’s (2018) conclusion that time-invariant amenities in Albouy (2016) can help us better understand housing price ‘dynamics’. Regarding the modelling of fixed effects, a more reasonable application of time-invariant indices can be seen in Favara & Imbs (2015) by investigating the joint effect of time-invariant and time-varying variables on housing returns. In this setting, the joint effect is not time-invariant and would not be absorbed by regional fixed effects.

Table 4. Housing market determinants

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
Constant	-0.004***	-0.004***	-0.003*	-0.003*	-0.004***	-0.006***	0.002	-0.030	0.010	-0.056***
	(0.001)	(0.0014)	(0.001)	(-0.001)	(0.001)	(0.001)	(0.003)	(0.004)	(0.008)	(0.012)
Income growth	0.618***	0.617***	0.618***	0.617***	0.618***	0.618***	0.784***	0.837***	0.596***	1.071***
	(0.024)	(0.024)	(0.024)	(0.024)	(0.024)	(0.024)	(0.038)	(0.055)	(0.026)	(0.068)
Population growth	0.502***	0.460***	0.479***	0.447***	0.478***	0.507***	0.029***	-0.136	1.406***	0.026**
	(0.065)	(0.007)	(0.065)	(0.066)	(0.065)	(0.065)	(0.093)	(0.130)	(0.073)	(0.127)
Quality		0.039***								
		(0.015)								
Productivity			0.034***							
			(0.008)							
Total amenity				0.036***						
				(0.008)						
Land regulation					0.005***					
					(0.001)					
Coast/lake						0.006***				
						(0.001)				
Supply elasticity							-0.004***			
							(0.001)			
Land share								0.069***		1.054***
								(0.008)		(0.018)
Time fixed									✓	✓
Regional fixed									✓	✓
Observations	5880	5880	5880	5880	5880	5880	2670	1260	5880	1260
Adjusted R^2	14.89%	14.98%	15.13%	15.14%	15.12%	15.08%	16.43%	22.43%	48.72%	56.67%

Notes: This table presents the result of estimating Equation (2). All of these analyses are of balanced panel data from 1986 to 2015. The number of markets can be reduced based on the availability of proposed data and indices. Numbers in parentheses denote the corresponding standard errors for the coefficients. ***, ** and * denote significance levels at 1%, 5% and 10%, respectively.

Table 5. Variance decomposition

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
Income growth	91.57%	92.24%	90.02%	90.98%	90.19%	89.65%	96.22%	74.93%	10.64%	20.18%
Population growth	8.43%	6.73%	7.47%	6.39%	7.43%	8.40%	0.02%	0.36%	7.54%	0.33%
Quality	·	1.03%	·	·	·	·	·	·	·	·
Productivity	·	·	2.51%	·	·	·	·	·	·	·
Total amenity	·	·	·	2.63%	·	·	·	·	·	·
Land regulation	·	·	·	·	2.38%	·	·	·	·	·
Coast/lake	·	·	·	·	·	1.96%	·	·	·	·
Supply elasticity	·	·	·	·	·	·	3.76%	·	·	·
Land share	·	·	·	·	·	·	·	24.71%	·	2.76%
Time fixed	·	·	·	·	·	·	·	·	70.97%	70.40%
Regional fixed	·	·	·	·	·	·	·	·	10.95%	6.34%
Observations	5880	5880	5880	5880	5880	5880	2670	1260	5880	1260
Adjusted R^2	14.89%	14.98%	15.13%	15.14%	15.12%	15.08%	16.43%	22.43%	48.72%	56.67%

Notes: Based on the result in Table 4, I perform the corresponding variance decomposition analysis. All of these analyses are of balanced panel data from 1986 to 2015. The number of markets can be reduced based on the availability of proposed data and indices. I first compute the partial sum of squares for each effect in the model and then normalise each estimate by the sum across the effects, forcing each column to sum to one.

5. Further Analysis

Sections 3 and 4 present results opposing the literature regarding the significance of national and time-invariant factors in housing markets. To further investigate this issue, I explore two-way fixed effects using different intercept dummy techniques and examine the prediction by controlling the time-series dimension.

5.1 Fixed Effects Model

I re-consider Equation (2), a linear unobserved effects model with two-way fixed effects for *i* regional markets and *t* time periods:

$$Housing\ Market_{i,t} = c + \gamma X_{i,t} + \alpha_i + \beta_t + \varepsilon_{i,t}, \tag{2}$$

where $X_{i,t}$ is a group of local factors, α_i is a regional fixed effect and β_t is a time fixed effect. A general way of modelling regional fixed effects involves using city dummies to allow the intercept to differ across regional markets, whereas the modelling of time fixed effects allows the intercept to vary across different time periods. In standard housing modelling, the unobserved effect of common macro factors at the national level can be absorbed by time fixed effects in a panel analysis, while the unobserved effect of time-invariant factors at the city level can be absorbed by regional fixed effects (Han, 2013; Favara & Imbs, 2015).

Technically, the estimated coefficient of an included time (city) dummy corresponding to a particular period (city) is an estimate of the difference between the intercept in that period (city) and the intercept in the omitted period (city). Accordingly, the time (β_t) and regional (α_i) fixed effects in Equation (2) can be expressed as:

$$\alpha_i = d_2\alpha_2 + d_3\alpha_3 + \cdots + d_i\alpha_i, \tag{3}$$

$$\beta_t = d_2\beta_2 + d_3\beta_3 + \cdots + d_t\beta_t, \tag{4}$$

where d is the dummy variable of 1 to indicate that period or city or 0 otherwise. Based on Equation (4), it can be seen that if one reduces the time series dimension, then the number of time dummies will be decreased. Thus, even though the coefficient of time series dummies (i.e., time fixed effects) are assumed to be significant in the modelling, its explanatory power in the modelling can be eliminated by reducing the time series dimension. An extreme case is that if the sample size is reduced to one period, then there exists no time fixed effects at all (i.e., 0% R-squared). Therefore, the application of time fixed effects typically requires a sample period of at least two periods.

In contrast, when the time series dimension in panel data decreases, regional fixed effects can be increasingly important. The underlying reason is that regional fixed effects wipe out explanatory variables that do not vary over time within a market. Therefore, with fewer time series observations, fewer time series variations will be identified within the market in panel data modelling. In the extreme case of a one period sample, all of the variation will be time-invariant. Therefore, a naive regional fixed effects model can lead to a 100% R-squared. This econometric effect potentially explains why the literature documents lower explanatory power for time-varying national factors (i.e., time fixed effects), yet higher explanatory power for time-invariant local factors (i.e., regional fixed effects) in housing markets.

5.2 Empirical Results

The sources of the data are consistent with Section 3—221 markets in a sample from 1985:Q2 to 2015:Q4. For simplicity, I start by considering time fixed effects alone in the modelling. Table 6 presents the explanatory power of time fixed effects in three different series: housing prices, returns and risk premiums based on different time series dimensions. It shows that when the time series dimension is expanded, the impact of time fixed effects will consistently increase, especially for housing returns and housing risk premiums.

Table 6. Adjusted R-squared of time fixed effects in housing

	1 quarter (1985:Q2)	10 quarters (1985:Q2 to 1987:Q3)	50 quarters (1985:Q2 to 1997:Q3)	100 quarters (1985:Q2 to 2010:Q1)	All Sample (1985:Q2 to 2015:Q4)
Price	NA (0%)	4.48%	53.49%	76.35%	74.10%
Return	NA (0%)	1.18%	6.23%	29.37%	31.70%
Risk premium	NA (0%)	1.91%	6.97%	29.73%	32.43%
Observations (quarters×markets)	221 (1×221)	2,210 (10×221)	11,050 (50×221)	22,100 (100×221)	27,183 (123×221)

Notes: I estimate $Housing\ Market_{i,t} = c + \beta_t + \varepsilon_{i,t}$ where $Housing\ Market_{i,t}$ is one of three time-series housing market variables: price, return (i.e., return calculated as the first difference of the natural logarithm) and risk premium (i.e., asset return minus the risk-free rate) for the dependent variable and β_t is a time fixed effect. Though the data are collected from 1985:Q1 to 2015:Q4, I request the sample to start at 1985:Q2 to allow three measures to have the same time-series dimension.

However, such a pattern does not hold for housing prices when the time series dimension expands from 100 quarters to the full sample of 123 quarters. One possible explanation for this finding is that national factors might not be important in some periods and might not explain the variation of housing price dynamics well in this circumstance. Thus, panel data modelling involving the use of some particular sample periods can weaken the overall significance of time fixed effects in housing markets. Despite this, when the sample size increases from 10 to 100 quarters in Table 6, the explanatory power of time fixed effects significantly increases from 4.48% to 76.35%. Though Glaeser et al. (2014) document that 8% of variation in housing prices is caused by year fixed effects, I further show that it is possible for us to reach this conclusion, especially by controlling the number of time series observations.

Next, I examine the prediction that the explanatory power of regional fixed effects will increase when the sample size is reduced, reporting the corresponding result in Table 7. Overall, the empirical evidence on housing prices, housing returns and housing risk premiums supports the prediction that regional fixed effects are important in panel data modelling, especially when the time series sample size is reduced. The underlying intuition is that when there is less time series variation in the sample, there will be less time-varying heterogeneity across regional markets in the panel data model for us to identify and control for. Therefore, the majority of variation in housing market dynamics will appear to be time-invariant.

Table 7. Adjusted R-squared of regional fixed effects in housing

	1 quarter (1985:Q2)	10 quarters (1985:Q2 to 1987:Q3)	50 quarters (1985:Q2 to 1997:Q3)	100 quarters (1985:Q2 to 2010:Q1)	All sample (1985:Q2 to 2015:Q4)
Price	NA (100%)	84.83%	18.57%	6.52%	8.78%
Return	NA (100%)	26.95%	2.27%	0.24%	0.24%
Risk premium	NA (100%)	26.68%	1.33%	0.23%	0.23%
Observations (quarters×markets)	221 (1×221)	2,210 (10×221)	11,050 (50×221)	22,100 (100×221)	27,183 (123×221)

Notes: I estimate $Housing\ Market_{i,t} = c + \alpha_i + \varepsilon_{i,t}$ where $Housing\ Market_{i,t}$ is one of three time-series housing market variables: price, return (i.e., return calculated as the first difference of the natural logarithm) and risk premium (i.e., asset return minus the risk-free rate) for the dependent variable and α_i is a regional fixed effect. Though the data are collected from 1985:Q1 to 2015:Q4, I request the sample to start at 1985:Q2 to allow three measures to have the same time-series dimension.

This can potentially explain why Beracha et al. (2018) find that time-invariant factors, such as local amenities in Albouy (2016), drive variation in house price 'dynamics'. Though Beracha et al. (2018) claim to collect an unbalanced panel dataset from 1975 to 2013, their empirical analysis focuses on the cross-sectional relationship between average housing returns and time-invariant amenities across MSAs. Without a time-series dimension, it is therefore no surprise that sub-regional fixed effects and time-invariant local amenities can explain 51.4% of variation in housing returns in Beracha et al. (2018). It should be noted that the concept of ‘dynamics’ requires empirical evidence over time, which is not incorporated in the empirical modelling of Beracha et al. (2018). In the extreme case of regional fixed effects for a one period sample, the regional fixed effects at the MSA level will identify that all of the factors are time-invariant within a market. This can lead to a naive result of explanatory power of 100% for explaining the variation in housing markets at the MSA level.

Table 8. Contribution of fixed effects in adjusted R-squared

	1 quarter (1985:Q2)	10 quarters (1985:Q2 to 1987:Q3)	50 quarters (1985:Q2 to 1997:Q3)	100 quarters (1985:Q2 to 2010:Q1)	All sample (1985:Q2 to 2015:Q4)
Panel A:					
Price	NA (100%)	76.39%	73.24%	83.68%	83.52%
Return	NA (100%)	11.27%	7.72%	29.91%	32.21%
Risk premium	NA (100%)	12.24%	8.45%	30.26%	32.92%
Panel B: Time fixed effects (regional fixed effects)					
Price	0%	26%	73%	91%	88.6%
	(NA/100%)	(74%)	(27%)	(9%)	(11.4%)
Return	0%	29%	67%	96%	96.8%
	(NA/100%)	(71%)	(33%)	(4%)	(3.2%)
Risk premium	0%	33%	69%	96%	96.9%
	(NA/100%)	(67%)	(31%)	(4%)	(3.1%)
Observations	221	2,210	11,050	22,100	27,183
(quarters×markets)	(1×221)	(10×221)	(50×221)	(100×221)	(123×221)

Notes: I estimate $Housing\ Market_{i,t} = c + \alpha_i + \beta_t + \varepsilon_{i,t}$ where $Housing\ Market_{i,t}$ is one of three time-series housing market variables: price, return (i.e., return calculated as the first difference of the natural logarithm) and risk premium (i.e., asset return minus the risk-free rate) for the dependent variable and α_i (β_t) is a regional (time) fixed effect. Though the data are collected from 1985:Q1 to 2015:Q4, I request the sample to start at 1985:Q2 to allow three measures to have the same time-series dimension. Panel A reports the adjusted R-squared, while Panel B reports the relative contribution of fixed effects to the adjusted R-squared.

Finally, I combine the results of time and regional fixed effects to further investigate whether the prediction of one-way fixed effects can robustly hold in the case of two-way fixed effects. Panel A in Table 8 reports the adjusted R-squared regarding the impact of time and regional fixed effects on housing market dynamics. Then, I decompose the variation in housing dynamics attributable to regional and time fixed effects, respectively. I compute the partial sum of squares for each effect in the model and then normalise each estimate by the sum across the effects, forcing each column to sum to one. The result in Panel B of Table 8 remains consistent with that of one-way fixed effects in Tables 6 and 7.

5.3 Implications

In this section, I provide an explanation for why time-varying national factors in the US housing markets are documented to be of little importance in Glaeser et al. (2014), and also why a high significance of time-invariant local factors is documented in Beracha et al. (2018). I argue that reducing the sample size of a time series can decrease the significance of time fixed effects, but increase the significance of regional fixed effects in panel data modelling of housing markets. Overall, empirical evidence from controlling sample sizes supports these two predictions. Therefore, without emphasising the time series dimension, I argue that the documented evidence for low or high explanatory power of national factors (i.e., time fixed effects) or time-invariant local factors (i.e. regional fixed effects) can be misleading.

This can be a serious issue in the modelling of housing market dynamics. Unlike financial markets, real estate price indices are not well established across countries. For instance, Wu, Gyourko and Deng (2012) publish the Wharton/National University of Singapore/Tsinghua Chinese Residential Land Price index from 2004 to the recent period with an annual frequency at the city level. Up to this period, there are only 14 time series observations at the city level (from 2004 to 2017) for China. In such a short time series sample, there is less time series variation within a market. Hence, I expect one can easily document low explanatory power for time fixed effects, but high explanatory power for city fixed effects, irrespective of the surveyed countries around the globe.

6. Conclusion

Though housing plays an important role in the economy, the evolution of housing price dynamics remains puzzling in the literature. While standard asset pricing models suggest a significant role of national factors in real estate markets, a series of recent literature of Hwang and Quigley (2006), Han (2013) and Glaeser et al. (2014) promotes a dominant role of local factors in housing markets as in line with the conventional urban models. Beracha et al. (2018) further suggest that time-invariant local amenities are critical drivers for housing price dynamics. This study seeks to better understand the nature of housing markets by empirically assessing the validity of the statements in Hwang and Quigley (2006), Han (2013), Glaeser et al. (2014) and Beracha et al. (2018).

Through variance decomposition analysis in the MSAs of the US over the period 1985 to 2015, I find that national factors, proxied by quarterly fixed effects, can account for around 74.42% of the variation in housing prices and 31.70% in housing returns. In contrast to the common presumption in the housing literature of Hwang and Quigley (2006), Han (2013) and Glaeser et al. (2014), I show that national factors play an important role in housing markets. Leung (2004) expresses concerns that some housing and urban economics research ignores the interplay between housing markets and the macroeconomy. My empirical evidence echoes Leung (2004) that theoretical and empirical analyses for urban and housing economics should include macroeconomic factors as control variables.

Furthermore, Beracha et al. (2018) conclude that the time-invariant local amenity in Albouy (2016) is a significant driver of house price 'dynamics' in the MSAs of the US. However, one limitation in Beracha et al. (2018) is the lack of time-series dimension for modelling housing price 'dynamics'. By incorporating a time-series dimension into a cross-sectional model, I show that time-invariant factors can be absorbed by regional fixed effects and contribute to explaining 8.78% of variation in housing prices and 0.24% of variation in housing returns. I further demonstrate that the significance of regional fixed effects can be artificially increased by reducing the time series dimension in panel data modelling. My results suggest that limited sample sizes in the time series dimension for modelling housing price dynamics can lead to the conclusion of Beracha et al. (2018).

Altogether, the findings from this research provide not only a better understanding of the nature of housing markets, but also a deeper insight as to implications for empirical and theoretical research regarding the assumptions of housing market dynamics and their interaction with the micro and macro economy.

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Paper 3. An Inflation Hedging Tale of Housing Markets

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Abstract

This paper investigates the long-run relationship between housing markets and macroeconomic fundamentals across 10 countries from 1871 to 2012, focusing on the inflation-hedging characteristics of residential real estate. Results show that inflation is the major source of understanding variation in housing price changes compared to other macroeconomic fundamentals (i.e., real income and population). However, widespread adoption of an inflation targeting policy from the early 1990s suggests a less prominent role for inflation-hedging thereafter. During this latter period, i.e. 1990 to 2012, much of the housing market variation is linked not to inflation risk but to changes in real income. The historical findings regarding inflation hedging property in housing markets have implications for portfolio management.

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1. Introduction

Inflation is one of the major risk sources in investment. Since the seminal work of Fisher (1930), inflation protection has received worldwide attention. Though inflation protection from homeownership has been studied in the literature, we know little about time-varying inflation hedging characteristics in relation to monetary environments. Since the 1990s, most countries have adopted the inflation targeting as a framework for monetary policy. An interesting question to address is whether the inflation protection from homeownership has changed. This paper aims to shed light on the time-varying inflation protection in housing markets across 10 countries in the long-run over 1871 to 2012. Using the inflation targeting regime as a natural setting, this study attempts to uncover the time-varying characteristics of inflation protection relative to other macroeconomic fundamentals for understating housing price dynamics.

Previous studies examining the relationship between housing prices and the macroeconomy are divided between economics and finance. The economic literature focuses on the role of macroeconomic variables, such as income, population and interest rates, in the determination of house prices (e.g., Adams & Fuss, 2010; Egert & Mihaljek, 2007; Englund & Ioannides, 1997; Holly & Jones, 1997; Holly, Pesaran & Yamagata, 2010; Hort, 1998; Kishor & Marfatia, 2017; Leung, 2014; Meen, 2002; Nneji, Brooks & Ward, 2013), while the finance literature (e.g., Bond & Seiler, 1998; Brounen, Eichholtz, Staetmans & Theebe, 2014; Fama & Schwert, 1997; Rubens, Bond & Webb, 1989) emphasises the hedging interplay between residential real estate and inflation.

This study combines and extends the two strands of the literature by investigating the fundamentals in housing assets from both an economic and financial perspective. I examine 10 countries (Australia, Denmark, Finland, France, the Netherlands, Japan, Norway, Sweden, Switzerland and the US) over a 100-year period, focusing on the implications of inflation protection from homeownership under different monetary regimes. This paper contributes to the existing literature in two ways. The first is by providing empirical evidence about the historical relationship between housing markets and macroeconomic fundamentals over the past century. Presenting a historical review of housing and macroeconomic fundamentals from 1871 to 2012 enables insight into the drivers of housing returns over the long term. This is of interest to investors and policy makers, especially since homeowners typically have long-term investment horizons. In the context of a historical review, this paper attempts to determine the extent to which macroeconomic variables explain the dispersion of housing price changes. Overall, I find that inflation is the stronger determinant of variation in housing prices, relative to population and real income.

The next question is whether this result always holds true, even when little inflation risk is present in the market for investors to hedge. Motivated by Fogler's (1984) argument that real estate acts as an effective inflation hedge during periods of high inflation, the second purpose of this paper is to address whether the dynamic relationship between housing markets and macroeconomic fundamentals has significantly changed since the inflation targeting framework implemented since the 1990s. Although the effect of inflation risk on asset prices is one of the long-standing questions in finance (e.g., Bond & Seiler, 1998; Brounen et al., 2014; Fama & Schwert, 1997; Rubens et al., 1989), the literature is silent on whether inflation hedging remains a critical driver of housing price dynamics in a period of inflation stabilisation. Using 1990 as a break point for two main sub-samples, I find economically and statistically significant evidence to support Fogler's (1984) conjecture because inflation risk explains only 0.21% of the variation in housing price changes from 1990 to 2012, in contrast to 12.73% from 1871 to 1989.

Following Kishor and Marfatia (2017), this study starts by examining the long-run relationship between macroeconomic factors and housing markets through conventional integration techniques. One of the underlying issues in identifying a long-run relationship is the lack of sufficiently long time series data. Therefore, a long-run relationship is typically defined as the stationary linear combination of nonstationary random variables. First, I consider conventional integration techniques, applying them to historical data from 1870 to 2012. My results show a long-run relationship between macroeconomic factors (e.g., real income, population and inflation) and housing prices in eight industrialised countries (Australia, Denmark, France, the Netherlands, Japan, Sweden, Switzerland and the US), suggesting that changes in house prices reflect movements in macro fundamentals in these countries.

A drawback of integration analysis is its limited scope for economic interpretation, especially regarding the relationships between macroeconomic fundamentals and housing prices. In line with the recent work of Eichholtz, Straetmans and Theebe (2012), Brounen et al. (2014) and Giglio, Maggiori and Stroebel (2015), this paper argues that the long-run effect of macroeconomic fundamentals can be better understood using standard regression analysis over long-term data. I estimate a hedonic pricing model with three macroeconomic fundamentals (e.g., real income, population and inflation) in each country over a sample period of 100 years. Historically, I find that more than half of the explained variation in this hedonic pricing model stems from inflation in 9 out of the 10 countries examined. This result suggests that inflation is the best candidate for understanding variations in housing price changes over the past century.

Next, I focus on empirical evidence since the 1990s, after which several countries have used numerical targets for inflation as a key component of monetary policy (e.g., Johnson, 2002). I hypothesise that less inflation risk exists during this period, obviating the need for investors to hedge against this factor and reducing the importance of inflation risk as a determinant of house price variation. Using panel data from 10 countries based on a sub-period from 1990 to 2012, I find that inflation risk alone accounts for only 0.21% of variation in housing price changes—a substantial drop from 12.22% over the entire sample. In contrast, real income contributes to the explanatory power of 26.13% in the 1990 to 2012 sub-sample but only 0.62% over the whole period. My result confirms that if inflation risk is not present in the market, there is little motivation for investors to hedge against inflation.

Among the extensive body of literature examining hedging against inflation in the residential real estate market, my paper is most related to Brounen et al. (2014), who re-examine inflation protection from homeownership in Amsterdam between 1814 and 2008, in a historical context. This paper extends the work of Brounen et al. (2014) to the international setting by providing international evidence for inflation protection from homeownership across 10 countries from 1871 to 2012. I find that inflation protection from homeownership is the major determinant of housing price dynamics in the setting of three fundamentals for a hedonic pricing model.

More broadly, I expand our understanding of inflation protection from homeownership under different monetary regimes. While most of the literature (e.g., Anari & Kolari, 2002; Brounen et al., 2014; Fama & Schwert, 1977; Hartzell, Heckman & Miles, 1987; Rubens et al., 1989; Stevenson, 2000) concurs that housing can serve as a hedge against inflation, this paper emphasises that this inflation hedging capacity can be masked in a period of inflation stabilisation. As hypothesised, there is lower inflation risk under an inflation targeting regime, and my empirical evidence supports the prediction that there has been limited inflation hedging benefit since the 1990s. Altogether, my results provide some of the first empirical insights into how the association between inflation and housing can vary in a historical context.

The remainder of the paper is structured as follows. Section 2 provides a literature review and motivation for this research. Sections 3 and 4 present the data description and the preliminary analysis, respectively. Section 5 reports the main results. Section 6 provides a comparative analysis of sub-samples. Section 7 discusses avenues for future research. Section 8 concludes.

2. Literature Review and Motivation

2.1 Literature Review

Changes in housing values play a critical role in household behaviour. Fluctuations in the value of housing alters, for instance, household balance sheets, perceived wealth and decisions about consumption and saving, which affect economic demand generally. Therefore, understanding the behaviour of housing markets has been of significant interest to academics, policymakers and economic agents.

A key concern of many economists and policymakers is whether macroeconomic fundamentals can explain changes in house prices. Assuming such a relationship exists in general, times during which housing prices move out of line with macroeconomic fundamentals indicate potential asset price bubbles, which might quickly turn into busts, resulting in economic contraction. Housing has a unique dual role as both a consumption and an investment good. Like financial assets, house prices can be highly volatile. According to the Standard & Poor's/ Case-Shiller US National Home Price Index, nominal housing rose by 65.5% between 2001 and 2006 and then fell by 23.10% between 2006 and 2011. Therefore, an understanding of the relationship between macroeconomic fundamentals and housing markets spans both economics and finance literatures.

Among the macroeconomic fundamentals cited as relevant to the housing market, the economic literature pays particular attention to income, population and interest rates. Income is a major contributor to housing affordability. Higher household income increases willingness to pay for housing, thus increasing housing demand and, at least theoretically, placing upward pressure on housing prices. Some empirical studies show a long-run relationship between housing prices and income (e.g., Holly & Jones, 1997; Hort, 1998; Kishor & Marfatia, 2017; Meen, 2002). Empirical research by Hwang and Quigley (2006), Egert and Mihaljek (2007) and Han (2013) further directly supports the theoretical prediction regarding the positive effect of income on housing prices.

Similar to the effect of household income, higher population growth also indicates a greater housing demand, and, therefore, greater house prices. Though most early studies find insignificant or negative effects of population growth on house prices (e.g., Hort, 1998), recent empirical work by Hwang and Quigley (2006) and Han (2013) supports the theoretical prediction of a positive effect of population growth on house prices in the US.

Since owning a house requires large capital outlay relative to household wealth, properties are typically financed through mortgages. Consequently, the other key factor affecting house prices is the rate of interest. Theoretically, a lower interest rate leads to a lower mortgage rate, therefore increasing housing demand and, in turn, leading to higher house prices. Nneji et al. (2013) show that short-term interest rates have an inverse relationship with housing prices in the 'boom', 'steady-state' and 'crash' regimes of US real estate markets. Based on international evidence from 15 countries over 30 years, Adams and Fuss (2010) provide empirical evidence for a negative association between long-term interest rates and house prices.

In contrast to the economic literature, the finance literature focuses on whether inflation is a key source of investment risk. The theory of inflation protection in asset prices goes back to Fisher (1930). According to Fisher (1930), the nominal rates of return on risky assets can be expressed as the sum of real returns and inflation, suggesting that inflation is an essential pricing factor in asset pricing models. However, across all asset classes, it is well established that, unlike stocks and bonds, housing has the ability to protect investors from the effects of inflation. If inflation is present in the market, investors would expect a resurgence of inflation to higher levels associated with higher levels of housing prices. The pioneering study by Fama and Schwert (1977) establishes that residential real estate serves as a complete hedge against both the expected and unexpected components of inflation (i.e., the difference between actual inflation and expected inflation); in contrast, common stocks cannot hedge against both expected and unexpected inflation, while government bonds and bills can only hedge against expected inflation.

Since then, a body of real estate finance literature has re-examined the hedging properties of private homes documented in Fama and Schwert (1977) by considering alternative definitions of expected and unexpected inflation. For instance, using expected inflation as estimated by the Livingston Survey, Rubens et al. (1989) show that residential real estate hedges against actual and unexpected inflation but not expected inflation in the US. Bond and Seiler (1998) differ by modelling the expected component of inflation using a first-order autoregressive moving average process and find that housing hedges against both expected and unexpected inflation.

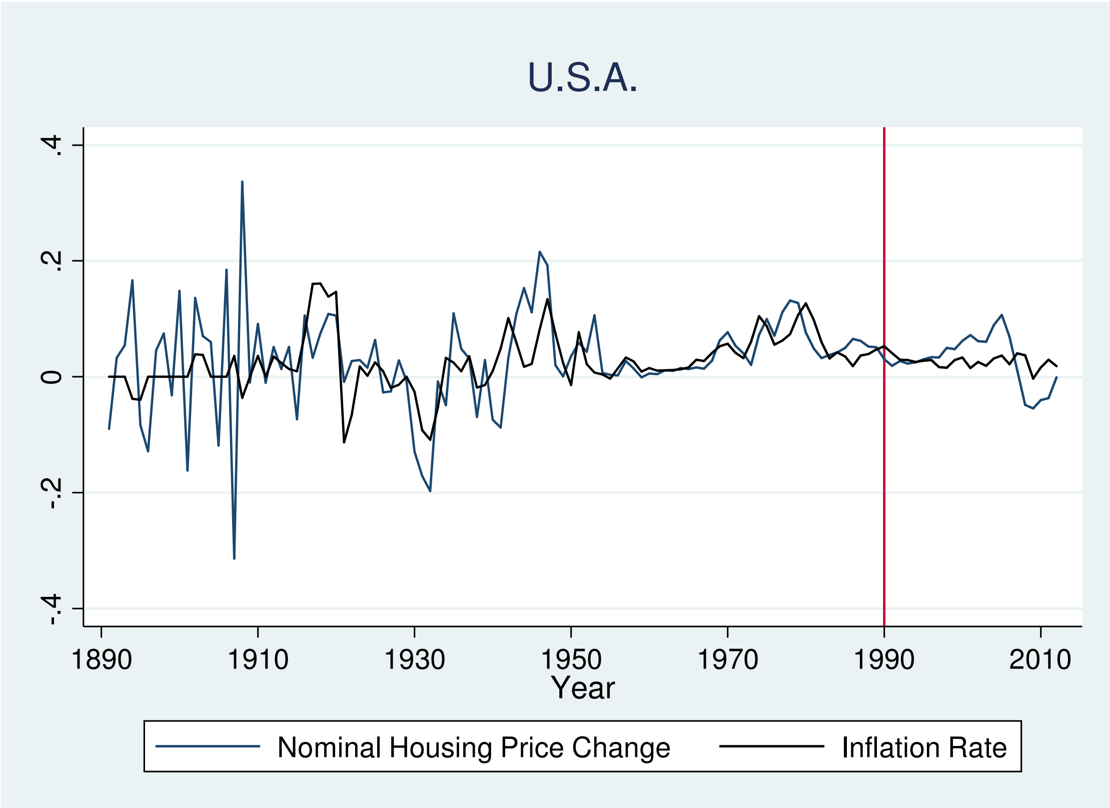
Through modelling the expected component of inflation using a first-order autoregressive time series model, recent research by Brounen et al. (2014) presents historical evidence for inflation protection from homeownership against actual and expected inflation in Amsterdam but weak evidence for hedging against unexpected inflation. As highlighted by Hartzell et al. (1987), there is substantial variation in the measurement of expected and unexpected inflation across the literature. This can help us understand the source of the contradiction regarding the effectiveness of hedging against expected and unexpected inflation. Collectively, the widespread consensus in the finance literature is that residential real estate acts as a hedge against actual inflation.

2.2 Motivation of This Research

As outlined previously, inflation protection has received considerable attention in the academic literature. Although the pioneering evidence of Fama and Schwert (1977) provides the foundation for the notion that real estate is an effective inflation hedge, the debate regarding the effectiveness of hedging against inflation remains inconclusive. Fogler (1984) emphasises that early empirical studies which highlight hedging effectiveness in real estate may rely heavily on the data employed during high inflation periods, as seen in Figure 1. Fogler (1984) argues that many countries exhibit strong performance from the property sector that are typically associated with sustained high rates of inflation. Therefore, Fogler (1984) argues that real estate serves as an effective inflation hedge, especially during long periods of high inflation.

Empirical evidence that directly investigates Fogler's (1984) stated association is scarce. To shed light on this issue, I argue that the inflation targeting policy adopted since 1990 provides a natural experiment for testing Fogler's (1984) assertion. Inflation targeting was pioneered in New Zealand in 1990, followed by Sweden and Australia in 1993 (Svensson, 2010). Using the US as an example, in Figure 1, I find that there is less inflation volatility especially since 1990, coinciding with less volatility in housing price changes. In contrast, there seems to be a high inflation risk before 1990, coupled with dramatic changes in housing prices. Clearly, the housing return–inflation relationship is pro-cyclical and sensitive to the monetary sector.

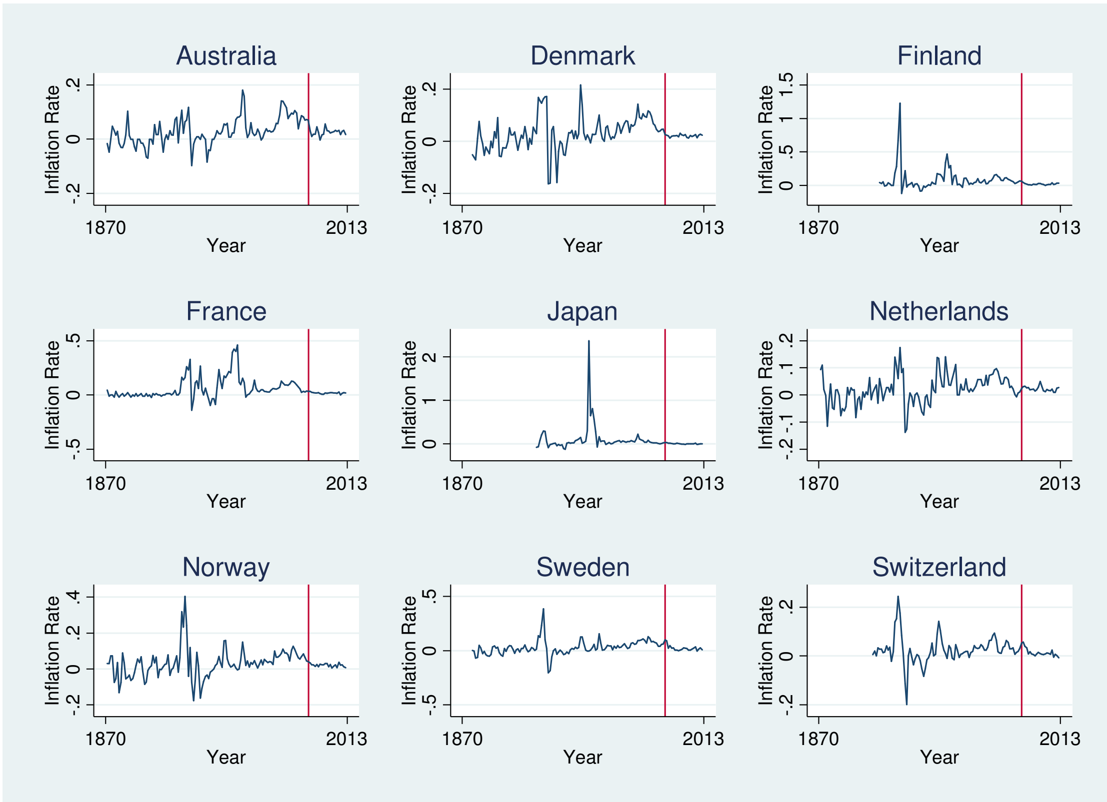
Figure 1. Inflation and housing price change movement in the US



Notes: I collect the nominal housing price and consumer price indices in Knoll, Schularick and Steger (2017) over the sample period from 1890 to 2012. Next, I construct the nominal housing price change and inflation rate by calculating the first difference of the natural logarithm of nominal housing price and consumer price indices. The vertical red line indicates the time period of 1990.

Though the Federal Reserve did not officially target inflation in 1990s, Mankiw (2001) highlights that the Federal Reserve discreetly implemented an inflation targeting policy. It was not until 2012 that the Federal Reserve set an explicit target inflation rate of approximately 2%, bringing the Federal Reserve in line with many of the world's other major central banks. To provide cross-country evidence of Fogler's (1984) conjecture, international markets outside the US are also considered. As highlighted in Figure 2, each of the nine additional countries examined seems to have experienced stable inflation since 1990. Hence, I hypothesise that if Fogler's (1984) argument—that real estate serves as a valid hedge in high inflation environments—holds, then the relative importance of inflation hedging in housing markets should diminish in a low inflation environment.

Figure 2. Inflation movement outside the US



Notes: The surveyed countries are outside the US. I collect the consumer price index in Knoll et al. (2017) over the sample period from 1870 to 2012. Then, I construct the inflation rate by calculating the first difference of the natural logarithm of the consumer price index. The vertical red line indicates the time period of 1990.

The contribution of this empirical exploration to the economic and finance literature is twofold. First, I provide a historical review comparing the relative importance of macroeconomic fundamentals in understanding housing price dynamics across international markets. Though the financial and economic fundamentals in housing markets have been examined respectively, a complete examination of all the factors in a historical context has not been carried out. Second, I compare the inflation hedging characteristic of residential real estate across inflation and non-inflation targeting regimes. While it is widely accepted that residential real estate is a hedging asset against actual inflation, little is known about whether this hedging characteristic can vary over time because of different monetary environments.

3. Summary Statistics

One of the fundamental issues associated with modelling the long-term relationship between housing markets and macroeconomic fundamentals is the lack of a sufficiently long time series of data. To investigate the long-run relationship between housing markets and the macroeconomy around the globe, I obtain an extensive dataset of the nominal housing price, real gross domestic product (GDP) per capita, population and consumer price index (CPI) from Knoll et al. (2017).[1] The database in Knoll et al. (2017) comprises annual data from 1870 to 2012 across 17 advanced economies and is collected from a broad range of historical sources and publications. I define the long-run relationship over 100 years and, thus, I require at least 100 years of uninterrupted data (i.e. without missing values) from an available sample period to 2012. After filtering out data which do not meet this criterion, my sample contains 10 countries: Australia,

[1] See data source in Knoll et al. (2017) for further details.

Denmark, Finland, France, the Netherlands, Japan, Norway, Sweden, Switzerland and the US.

Table 1 reports the summary statistics of housing price growth and inflation. Housing price growth and the rate of inflation are calculated as the first difference of the natural logarithm of the nominal housing price index and CPI, respectively. The nominal housing price growth is higher than the inflation rate for all countries. The substantially greater average inflation rates in Japan and Finland are mainly driven by outliers resulting from unique historical events. In particular, after World War II, Japan was a defeated country with an inflation rate of 238% in 1945. Finland declared independence in 1917, which was followed by a civil war in 1918, leading to an inflation rate of 123% in that year.

Table 1. Summary statistics for housing price growth and inflation

	Mean (%)	SD (%)	Min (%)	Max (%)	Obs
Panel A: Housing return					
Australia (1871 to 2012)	4.63	10.11	-20.47	84.38	142
Netherlands (1871 to 2012)	3.20	9.15	-34.52	32.53	142
Switzerland (1902 to 2012)	3.13	5.61	-10.50	21.38	111
Denmark (1876 to 2012)	3.91	7.55	-17.52	23.68	137
Finland (1906 to 2012)	9.15	17.16	-24.36	89.84	107
France (1871 to 2012)	6.65	8.28	-12.10	40.05	142
Norway (1871 to 2012)	4.18	8.46	-18.27	37.78	142
Sweden (1876 to 2012)	3.65	7.60	-37.47	21.84	137
Japan (1914 to 2012)	10.81	20.77	-24.85	98.06	99
US (1891 to 2012)	3.09	8.15	-31.39	33.69	122
Panel B: Inflation					
Australia (1871 to 2012)	2.94	4.58	-9.70	18.10	142
Netherlands (1871 to 2012)	2.24	4.99	-13.80	17.49	142
Switzerland (1902 to 2012)	2.33	4.99	-19.85	24.38	111
Denmark (1876 to 2012)	2.92	5.89	-16.36	21.57	137
Finland (1906 to 2012)	7.35	15.22	-12.04	122.79	107
France (1871 to 2012)	5.47	9.41	-14.06	46.14	142
Norway (1871 to 2012)	3.01	7.31	-17.63	40.42	142
Sweden (1876 to 2012)	2.83	6.17	-20.39	38.49	137
Japan (1914 to 2012)	8.01	26.89	-12.25	237.55	99
US (1891 to 2012)	2.67	4.49	-11.28	16.08	122

Notes: The data are collected from Knoll et al. (2017). The housing return and inflation is calculated as the first difference of the natural logarithm of the nominal housing price index and CPI, respectively.

Table 2 reports the summary statistics of real income and population growth. In terms of national income, Japan has the highest real income growth per year (2.83%), while Switzerland has the lowest (1.41%). According to the Migration Policy Institute, Australia and the US are often described as traditional immigration countries.[2] Australia and the US demonstrate the highest population growth in part because of the significant growth in immigration over the past century.

Table 2. Summary statistics for real income and population growth

	Mean (%)	SD (%)	Min (%)	Max (%)	Obs
Panel A: Real income growth					
Australia (1871 to 2012)	1.46	4.15	-15.94	13.10	142
Netherlands (1871 to 2012)	1.56	6.82	-41.08	51.37	142
Switzerland (1902 to 2012)	1.41	3.57	-11.20	11.88	111
Denmark (1876 to 2012)	1.71	3.75	-15.94	13.31	137
Finland (1906 to 2012)	2.41	4.79	-18.21	19.03	107
France (1871 to 2012)	1.67	6.26	-26.72	28.89	142
Norway (1871 to 2012)	2.13	3.56	-11.60	14.98	142
Sweden (1876 to 2012)	2.04	3.46	-13.60	10.35	137
Japan (1914 to 2012)	2.83	6.70	-37.24	14.05	99
US (1891 to 2012)	1.96	5.03	-14.67	16.06	122
Panel B: Population growth					
Australia (1871 to 2012)	1.79	1.07	-5.80	4.36	142
Netherlands (1871 to 2012)	1.09	0.51	-1.27	3.89	142
Switzerland (1902 to 2012)	0.78	0.63	-1.11	2.82	111
Denmark (1876 to 2012)	0.76	0.38	-0.07	1.41	137
Finland (1906 to 2012)	0.63	0.40	-0.61	1.55	107
France (1871 to 2012)	0.37	0.73	-3.47	2.47	142
Norway (1871 to 2012)	0.73	0.32	-1.56	1.48	142
Sweden (1876 to 2012)	0.55	0.30	-0.04	1.26	137
Japan (1914 to 2012)	0.91	0.61	-1.24	2.57	99
US (1891 to 2012)	1.31	0.43	0.48	2.09	122

Notes: The data are collected from Knoll et al. (2017). The growth of real income and population is calculated as the first difference of the natural logarithm of real GDP per capita and population, respectively.

4. Preliminary Analysis

In this section, I provide a preliminary examination of the long-run relationship between macroeconomic fundamentals and housing markets, using the standard cointegration technique. I first examine whether the group of nonstationary variables can

[2] See Migration Policy Institute (2018) for further details.

be integrated into the same order. If this condition is satisfied, I next examine whether these variables share a common trend.

4.1 Unit Root Test

Unlike the standard empirical literature investigating the theoretical prediction about macroeconomic effects on housing markets, some literature examines the long-run relationship between macroeconomic fundamentals and housing markets using integration techniques. While most studies support a long-run relationship (Adams & Fuss, 2010; Egert & Mihaljek, 2007; Holly & Jones, 1997; Holly et al., 2010; Hort, 1998; Kishor & Marfatia, 2017; Leung, 2014; Meen, 2002), other studies (Gallin, 2006; Mikhed & Zemcik, 2009) find opposing evidence. Motivated by this debate, this empirical research starts by investigating whether housing markets share a long-run relationship with macroeconomic fundamentals (e.g., population, real income and inflation).

I begin by formally testing for the presence of unit roots in house price index, real income, population and CPI data through the augmented Dickey–Fuller test (Dickey & Fuller, 1979):

$$\Delta y_t = \alpha + \beta y_{t-1} + \delta_t + \tau_1 \Delta y_{t-1} + \tau_2 \Delta y_{t-2} + \cdots + \tau_k \Delta y_{t-k} + \epsilon_t, \tag{1}$$

where y_t is the index at time *t*, δ_t represents the time trend and k is the number of lags determined by Bayesian information criterion (BIC). The null hypothesis of the augmented Dickey–Fuller test is that the variable is nonstationary (i.e., a unit root process with $\beta = 0$). If the null hypothesis is rejected, then the variable is suggested to be stationary.

Table 3. Unit root test

	Housing price	Population	Income	Consumer price
Panel A: Level series				
Australia (1870 to 2012)	1.95	0.87	0.34	1.29
Netherlands (1870 to 2012)	-2.12	-2.15	-0.76	0.71
Switzerland (1901 to 2012)	-0.31	-2.20	-2.10	-1.34
Denmark (1875 to 2012)	-0.05	-1.29	-1.55	0.02
Finland (1905 to 2012)	0.76	-1.24	-1.78	-0.71
France (1870 to 2012)	0.43	-0.86	-1.08	-0.80
Norway (1870 to 2012)	5.18	-1.41	-0.60	-0.72
Sweden (1875 to 2012)	3.46	-2.76	-0.36	-0.57
Japan (1913 to 2012)	-1.57	1.09	-1.87	-1.99
US (1890 to 2012)	0.06	-1.31	-1.09	0.80
Panel B: First difference				
Australia (1871 to 2012)	-7.24***	-4.05***	-6.34***	-5.05***
Netherlands (1871 to 2012)	-4.93***	-2.94**	-6.50***	-5.73***
Switzerland (1902 to 2012)	-6.27***	-3.09**	-5.52***	-5.52***
Denmark (1876 to 2012)	-6.47***	-1.99**	-10.90***	-3.67***
Finland (1906 to 2012)	-5.58***	-3.96***	-4.80***	-5.54***
France (1871 to 2012)	-4.42***	-4.76***	-10.79***	-4.07***
Norway (1871 to 2012)	-11.10***	-3.89***	-10.55***	-5.62***
Sweden (1876 to 2012)	-7.68***	-4.23***	-12.60***	-5.73***
Japan (1914 to 2012)	-3.46**	-2.35**	-5.81***	-4.48***
US (1891 to 2012)	-5.13***	-2.02**	-6.54***	-5.52***

Notes: This table reports the test statistics of the augmented Dickey–Fuller test for stationarity. ***, ** and * denote significance levels at 1%, 5% and 10%, respectively.

Investigating the stationarity of these series is essential since the validity of empirical analysis via the ordinary least squares approach hinges upon the stationarity assumption. If the repressors and regressands are nonstationary, the regression can generate spurious results. Table 3 reports the stationarity of the time series data. I find that the null hypothesis of non-stationarity at the level series (housing, population, real income and CPI) cannot be rejected. This indicates that the index is nonstationary at the level series. Therefore, I take the first difference of the natural logarithm of the level series. Based on the first difference series, I find that the null hypothesis of non-stationarity is rejected at 5% significance for all series. This shows that the first difference series is stationary and suitable for standard ordinary least square analysis.

4.2 Engle-Granger Testing for Integration

Since the nonstationary variables are integrated into the same order, it is possible that these nonstationary variables share a common trend in the long run. Following Kishor and Marfatia (2017), I employ the two-step cointegration test by Engle and Granger (1987) to investigate the long-run relationship between housing prices and macroeconomic environments by specifying and estimating the following long-run cointegrating equilibrium relationship:

$$H_t = \alpha + C_1 RI_t + C_2 P_t + C_3 C_t + \varepsilon_t, \tag{2}$$

where H_t, RI_t, P_t and C_t represent housing price, real income, population and CPI, respectively. If the variables in the proposed system share a common trend, the estimated integrating residuals ($\hat{\varepsilon}_t = H_t - \hat{\alpha} - \hat{C}_1 RI_t - \hat{C}_2 P_t - \hat{C}_3 C_t$) would be stationary.

Table 4. Engle–Granger testing for integration

	Test statistics
Australia (1870 to 2012)	-2.46***
Netherlands (1870 to 2012)	-2.78***
Switzerland (1901 to 2012)	-2.11**
Denmark (1875 to 2012)	-3.11***
Finland (1905 to 2012)	-0.89
France (1870 to 2012)	-3.36***
Norway (1870 to 2012)	-0.98
Sweden (1875 to 2012)	-2.42***
Japan (1913 to 2012)	-1.88**
US (1890 to 2012)	-3.82***

Notes: This table reports the test statistics of Engle–Granger testing for integration. ***, ** and * denote significance levels at 1%, 5% and 10%, respectively.

Table 4 summarises the results for stationarity of the cointegrating residuals. I find that the null hypothesis of unit root in the estimated residuals is rejected at 5% significance except for Norway and Finland.[3] This implies that, for most countries, there

[3] The exception of Finland and Norway implies that housing markets do not share a common trend with the macroeconomic environment of the three-factor setting, suggesting deviations of market prices away from these market fundamental values.

is a long-run relationship between macroeconomic fundamentals (i.e., real income, population and CPI) and housing markets.

Overall, the stationarity of estimated residuals in standard Engle–Granger testing for integration indicates consistent results for most countries. In line with previous literature (Egert & Mihaljek, 2007; Holly & Jones, 1997; Hort, 1998; Kishor & Marfatia, 2017; Meen, 2002), the empirical evidence supports a long-run relationship between macroeconomic fundamentals and the housing markets of Australia, Denmark, France, the Netherlands, Japan, Sweden, Switzerland and the US. The results imply that macroeconomic fundamentals and housing markets may drift apart temporarily in these countries, yet their tendency is to return to their long-run equilibrium.

5. Empirical Analysis

One problem with long-run equilibrium analyses using integration techniques is their limited scope for economic interpretation. This section employs the standard ordinary least square regression to explore the theoretical relationship between three macroeconomic factors (i.e., real income, population and inflation) and housing markets over a long-term sample period of 1871 to 2012.

5.1. Housing and Fundamentals

While housing research[4] typically applies the integration method over a short period to draw inferences about long-run relationships, recent work by Eichholtz et al.

[4] See Holly and Jones (1997), Hort (1998), Meen (2002), Gallin (2006), Egert and Mihaljek, (2007), Mikhed and Zemcik (2009), Adams and Fuss (2010), Holly et al. (2010), Leung (2014) and Kishor and Marfatia (2017).

(2012), Brounen et al. (2014) and Giglio et al. (2015) defines long-run evidence as empirical evidence over long-term time series data. The definition of the latter is more intuitive and more associated with housing, given its long investment horizon. Accordingly, I examine the long-term association between housing and macroeconomic fundamentals by estimating the following regression for each country based on unique, long-term data from 1871 to 2012:

$$HC_t = \alpha + C_1 GRI_t + C_2 PG_t + C_3 I_t + \varepsilon_t, \tag{3}$$

where HC_t is the housing price change, GRI_t is the growth of real income, PG_t is the population growth and I_t is the inflation. The income and population variables account for major economic and demographic factors in housing markets and are thus incorporated to control for the consumption demand for housing (e.g., Han, 2013). Residential real estate is known to provide inflation protection and is thus included to assess the inflation hedging efficiency of housing (e.g., Fama & Schwert, 1977). Finally, after estimating the baseline model, I decompose the variation in housing returns attributable to different factors by applying a parametric framework, analysis of covariance (ANCOVA).

Table 5 presents the empirical results from testing the relationship between housing price changes and the identified fundamentals, as per Equation (3). Intuitively, an increase in household income should spur housing demand and thereby increase housing prices. In support of this prediction, the empirical evidence in Table 5 shows a significantly positive relationship between income and housing in Australia, Denmark, Finland and Norway at the 5% significance level, although the relationship is insignificant in other countries. In sum, the results in Table 5 suggest a heterogeneous relationship between housing and real income across countries.

Theoretically, greater population growth can lead to higher housing demand, thereby increasing housing prices. As expected, I find a significantly positive relationship between population growth and housing prices in Australia, Finland, France and Japan at the 5% significance level. The insignificant or negative effects of population growth on house prices in Sweden is consistent with Hort's (1998) result regarding Sweden from 1970 to 1994.

Table 5. Fundamentals and housing

Country	Constant	Population	Income	Inflation	Adjusted R-squared	Obs
Australia	-0.020	1.745**	0.530***	0.905***	24.44%	142
	(0.015)	(0.690)	(0.179)	(0.162)		
Switzerland	0.021**	0.130	0.171	0.305***	4.19%	111
	(0.01)	(0.854)	(0.156)	(0.110)		
Denmark	0.012	-0.572	0.666***	0.703***	30.96%	137
	(0.014)	(1.454)	(0.148)	(0.097)		
Finland	-0.001	7.997**	1.101***	0.326***	9.19%	107
	(0.034)	(4.014)	(0.381)	(0.121)		
France	0.03***	3.638***	-0.062	0.372***	24.03%	142
	(0.007)	(0.940)	(0.113)	(0.067)		
Japan	-0.065**	13.92***	0.401	0.447***	37.07%	99
	(0.032)	(2.846)	(0.280)	(0.069)		
Netherlands	0.035**	-1.677	-0.061	0.711***	15.15%	142
	(0.017)	1.396	0.104	0.143		
Norway	0.030	-0.872	0.415**	0.294***	6.53%	142
	(0.018)	(2.171)	(0.197)	(0.095)		
Sweden	0.039	-2.014	0.074	0.263**	3.28%	137
	0.015	(2.140)	0.189	0.107		
US	-0.010	1.818	-0.108	0.730***	13.10%	122
	0.0240	1.656	0.139	(0.159)		

Notes: This table reports the relationship between macroeconomic fundamentals and housing markets across countries based on estimating Equation (3). ***, ** and * denote significance levels at 1%, 5% and 10%, respectively.

The most striking result in Table 5 is the significantly positive relationship between housing and inflation across all 10 countries, indicating that residential real estate has, historically, served as a hedge against inflation. Overall, in contrast to the inconsistent results regarding the relationship between non-inflation fundamentals (i.e., real income and population) and housing, the relationship between housing and inflation is significantly positive across the 10 countries.

To identify the dominant driver of housing price changes, I further decompose the sources of explanatory power (i.e., the model fitness). In the three-factor fundamentals model, Table 6 shows that inflation plays the most important role in housing markets for all countries except Finland. The cross-country comparison indicates that at least 61.97% of the explained variation can be attributed to inflation. In the case of Finland, inflation remains important, accounting for 37.18% of the explained variation. Altogether, among the proposed economic fundamentals, the historical evidence suggests that inflation is the most influential factor in housing price dynamics.

Table 6. Variance decomposition

Country	Adjusted R-squared	Variance decomposition			Obs
		Population	Income	Inflation	
Australia	24.44%	13.78%	18.95%	67.27%	142
Switzerland	4.19%	0.26%	13.41%	86.33%	111
Denmark	30.96%	0.21%	27.53%	72.26%	137
Finland	9.19%	20.22%	42.60%	37.18%	107
France	24.03%	32.48%	0.64%	66.88%	142
Japan	37.07%	35.03%	3.00%	61.97%	99
Netherlands	15.15%	5.43%	1.30%	93.27%	142
Norway	6.53%	1.14%	31.47%	67.39%	142
Sweden	3.28%	12.50%	2.14%	85.36%	137
US	13.10%	5.24%	2.64%	92.12%	122

Notes: Based on the result in Table 5, I perform the corresponding variance decomposition analysis. I first compute the partial sum of squares for each effect in the model and then normalise each estimate by the sum across the effects, forcing each row for variance decomposition to sum to one.

5.2 Expected and Unexpected Inflation

In addition to shedding light on whether residential real estate offers protection against actual inflation (i.e., observed inflation or ex-post realised inflation), I also quantify the relation between actual returns and expected inflation because this comes closer to Fisher's (1930) hypothesis. Fama and Schwert (1977) establish that realised inflation can be broken into expected and unexpected components, which lead to the following regression equation:

$$R_t = \alpha + C_1 E_{t-1}(\pi_t) + C_2[\pi_t - E_{t-1}(\pi_t)] + \varepsilon_t, \tag{4}$$

where R_t is the asset returns, $E_{t-1}(\pi_t)$ represents the expected inflation; unexpected inflation $[\pi_t - E_{t-1}(\pi_t)]$ is defined as the actual inflation minus the expected inflation. Fama and Schwert (1977) further argue that a reasonable proxy for expected inflation should satisfy the following condition:

$$\pi_t = \alpha + C_1(\beta_t) + \varepsilon_t, \tag{5}$$

where π_t represents the inflation rate and β_t is the proxy for expected inflation. The proposition of C_1 is close to one (i.e., a one to one correspondence) if β_t is a reasonable proxy. The empirical evidence of Fama and Schwert (1977) shows that the one month treasury bill serves as a proxy for the expected inflation rate.

Motivated by this, I augment the dataset of Knoll et al. (2017) by adding the short-term interest rate from the Jorda–Schularick–Taylor Macrohistory Database (Jorda, Schularick, & Taylor, 2016).[5] I consider that the historical time series dataset of short-term interest is a possible proxy for the expected inflation rate. Table 7 reports the summary statistics for short-term interest rates, showing that the average interest rate differs greatly across countries from 2.76% in Switzerland to 6.53% in Finland.

[5] See Jorda et al. (2016) for data sources and constructions. I splice the missing data with an alternative interest rate source from the Jorda–Schularick–Taylor Macrohistory Database in Jorda et al. (2016).

Table 7. Summary statistics for short-term interest rates

	Mean (%)	SD (%)	Min (%)	Max (%)	Obs
Australia (1871 to 2012)	5.06	2.98	1.25	15.18	142
Netherlands (1871 to 2012)	3.51	2.15	0.15	11.01	142
Switzerland (1902 to 2012)	2.76	1.83	-0.02	9.69	111
Denmark (1876 to 2012)	5.90	3.00	0.08	16.92	137
Finland (1906 to 2012)	6.53	3.15	0.22	16.50	107
France (1871 to 2012)	4.25	2.93	0.05	15.30	142
Norway (1871 to 2012)	5.46	2.83	1.80	14.61	142
Sweden (1876 to 2012)	5.09	2.56	0.40	14.49	137
Japan (1914 to 2012)	5.53	3.21	0.00	12.54	99
US (1891 to 2012)	4.09	2.93	0.10	16.39	122

Note: The short-term interest rate is collected from the Jorda–Schularick–Taylor Macrohistory Database in Jorda et al. (2016).

Using the short-term interest rate from Jorda et al. (2016) and inflation from Knoll et al. (2017), I estimate Equation (5) and report the results in Table 8. The coefficient estimate for inflation on housing is not close to one across the 10 countries. The adjusted R-squared is also low, ranging from -0.94% to 13.51%, compared to the documented model fitness of 82% based on semi-annual statistics in Fama and Schwert (1977).

Table 8. CPI and short-term interest rate

Country	Constant	CPI	Adjusted R-squared	Obs
Australia	0.002	0.532***	11.35%	142
	(0.007)	(0.122)		
Netherlands	0.009	0.390**	2.12%	142
	(0.008)	(0.193)		
Switzerland	0.006	0.605**	4.10%	111
	(0.008)	(0.25)		
Denmark	-0.012	0.695***	11.93%	137
	(0.01)	(0.158)		
Finland	0.077**	-0.048	-0.94%	107
	(0.034)	(0.471)		
France	0.041***	0.336	0.39%	142
	(0.014)	(0.269)		
Norway	0.004	0.474**	2.71%	142
	(0.013)	(0.214)		
Sweden	-0.009	0.724***	8.39%	137
	(0.011)	(0.198)		
Japan	0.094*	-0.251	-0.94%	99
	(0.054)	(0.851)		
US	0.003	0.578***	13.51%	122
	(0.007)	(0.130)		

Note: This table reports the empirical estimation of Equation (5): $\pi_t = \alpha + C_1(\beta_t) + \varepsilon_t$ where π_t is the inflation rate and β_t is the short-term interest rate.

The results indicate that using the short-term interest rate in Jorda et al. (2016) as a proxy for the expected inflation rate is unsatisfactory. Jorda et al. (2016) patch together the historical short-term interest rate from different resources. Therefore, the definition of the short-term interest rate is inconsistent. The inconsistent sources of historical data pose challenges for modelling expected inflation in the long-run.

Another strand of literature (e.g., Bond & Seiler, 1998; Brounen et al., 2014) assumes that actual inflation follows a first-order autoregressive (AR(1)) process and uses this process to decompose inflation into its expected and unexpected components. This approach is built on the intuition that inflation series are persistent but stationary and predictable based on the previous period's inflation (π_{t-1}):

$$\pi_t = \alpha + C_1(\pi_{t-1}) + \varepsilon_t. \tag{6}$$

If the AR(1) process is a reasonable method for modelling expected inflation, it should be noted that identifying whether C_1 is close to one is a prerequisite condition in theory, as is Equation (5), since it indicates a one-to-one relationship between observed and expected inflation. This condition is rarely checked in the literature before application. I thus investigate this issue. Table 9 reports the results of estimating Equation (6) for the 10 countries. The coefficient estimates of returns from the last period are not close to 1, ranging from 0.428 to 0.712. This result indicates that using AR(1) to model expected inflation is also not satisfactory.

Table 9. AR(1) model

Country	Constant	Lagged returns	Adjusted R-squared	Obs
Australia	0.009***	0.690***	47.49%	142
	(0.003)	(0.061)		
Netherlands	0.010**	0.541***	29.15%	142
	(0.004)	(0.071)		
Switzerland	0.007*	0.702***	48.75%	111
	(0.004)	(0.0686)		
Denmark	0.012***	0.589***	34.66%	137
	(0.005)	(0.069)		
Finland	0.042***	0.428***	17.51%	107
	(0.015)	(0.089)		
France	0.015**	0.712***	50.28%	142
	(0.006)	(0.006)		
Norway	0.012**	0.582***	33.46%	142
	(0.005)	(0.069)		
Sweden	0.009**	0.668***	44.27%	137
	(0.004)	(0.064)		
Japan	0.046*	0.437***	18.34%	99
	(0.026)	(0.092)		
US	0.01***	0.632***	39.61%	122
	(0.004)	(0.070)		

Note: This table reports the empirical estimation of Equation (6): $\pi_t = \alpha + C_1(\pi_{t-1}) + \varepsilon_t$ where π_t is the inflation rate at time *t* and π_{t-1} is the inflation rate at time *t-1*.

This empirical exercise demonstrates the challenges of finding a suitable proxy for expected inflation. It helps explain why Fama and Schwert (1977) find inflation hedging benefits against both expected and unexpected inflation, while Brounen et al. (2014) only find evidence for benefits against expected inflation. This exercise also reinforces the argument of Hartzell et al. (1987) that the observed ability of real estate to hedge against inflation depends on the method of measuring expected and unexpected inflation. Given the unavailability of a reasonable proxy for the treasury bill rate in a historical context, this paper focuses on the relationship between actual asset returns and actual inflation, albeit with an emphasis on the effect of monetary environments.

5.3 Panel Data Modelling

Since my dataset contains both time series and cross-sectional dimensions, I further expand Equation (3) by incorporating fixed effects for panel data modelling:

$$HC_{i,t} = \alpha + C_1 GRI_{i,t} + C_2 PG_{i,t} + C_3 I_{i,t} + \alpha_i + \beta_t + \varepsilon_{i,t}, \quad (7)$$

where $HC_{i,t}$ is the housing price change, $GRI_{i,t}$ is the growth of real income, $PG_{i,t}$ is the population growth, $I_{i,t}$ is the inflation, α_i is a regional fixed effect and β_t is a time fixed effect. In a panel analysis, the unobserved effect of common macro factors at the international level is absorbed by time fixed effects, while country fixed effects capture permanent differences across countries, such as those pertaining to geographical constraints and national amenities.

Table 10 reports the results of the panel data regression modelling. When inflation alone is included in the regression, inflation can account for 12.22% of variation in housing price changes. In the setting of three macroeconomic fundamentals, inflation risk contributes to 84.69% of the explained variation. This evidence further shows that inflation has a greater influence than real income and population on house prices globally.

However, in Column (7), the role of inflation becomes less prominent because some of the effect of inflation risk is absorbed by time fixed effects. In the five macroeconomic fundamentals with the explanatory power of 20.63%, time fixed effects contribute to 60.87% of the explained variation, while inflation contributes to 18.74%. From the modelling expansion from Columns (4) to (7), it is interesting to note that the dominant role of inflation becomes minor once time fixed effects are incorporated. This indicates that inflation risk in asset prices can be highly linked to time fixed effects or systematic risk at the international level. Collectively, inflation remains most essential in housing markets among the proposed three observable macroeconomic fundamentals.

Table 10. Panel data modelling from 1871 to 2012

	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Panel A: Regression modelling							
Constant	0.047*** (0.003)	0.037*** (0.003)	0.042*** (0.005)	0.018*** (0.005)	-0.009 (0.012)	-0.0003 (0.05)	0.005 (0.050)
Income	0.185*** (0.062)			0.306*** (0.058)	0.262*** (0.058)	0.101 (0.068)	0.046 (0.067)
Inflation		0.365*** (0.027)		0.397*** (0.027)	0.368*** (0.027)	0.304*** (0.033)	0.260*** (0.033)
Population			0.890** (0.430)	1.287*** (0.401)	2.249*** (0.481)	1.089** (0.444)	2.371*** (0.572)
Country fixed effects					✓		✓
Time fixed effects						✓	✓
Adjusted R-squared	0.62%	12.22%	0.26%	14.80%	17.11%	17.88%	20.63%
Observations	1, 281	1, 281	1, 281	1, 281	1, 281	1, 281	1, 281
Panel B: Variance decomposition							
Income	1	·	·	11.17%	7.70%	0.80%	0.15%
Inflation	·	1	·	84.69%	67.32%	29.68%	18.74%
Population	·	·	1	4.14%	8.21%	2.14%	5.31%
Country fixed effects	·	·	·	·	16.77%	·	14.94%
Time fixed effects	·	·	·	·	·	67.38%	60.87%
Adjusted R-squared	0.62%	12.22%	0.26%	14.80%	17.11%	17.88%	20.63%

Notes: Panel A in Table 10 presents the results of estimating Equation (7). ***, ** and * denote significance levels at 1%, 5% and 10%, respectively. Based on the result in Panel A, I first compute the partial sum of squares for each effect in the model and then normalise each estimate by the sum across the effects, forcing each column for variance decomposition to sum to one in Panel B.

6. Comparison: Inflation Targeting

In Section 5, I document the important role played by inflation in housing markets during the period from 1871 to 2012. In this section, I explore whether the dominance of inflation hedging in housing holds in different sub-periods by employing the inflation targeting regime as a threshold.

6.1 Inflation Targeting

One of the main objectives of central banks across the world is to achieve price stability, preferably with moderate inflation rather than no inflation or deflation. According to Mankiw (2001), stabilisation of inflation is desirable for several reasons, two of which are particularly relevant to macroeconomics and finance. First, inflation makes currency less reliable as a yardstick for measuring value. Second, highly volatile inflation can generate unnecessary risk between debtors and creditors, due to unexpected changes in price levels.

Since the 1990s, many countries have adopted monetary policies that focus on maintaining a low level of inflation. Inflation targeting was first introduced by New Zealand in 1990, followed by Sweden and Australia in 1993 (Svensson, 2010). Though few are officially acknowledged as inflation targeters, most countries across the globe appear to stabilise inflation risk, as evidenced in Figures 1 and 2 of Section 2. To demonstrate, though prior to 2012 the Federal Reserve did not officially target inflation in the US, Mankiw (2001) indicates that the Federal Reserve has discreetly implemented an inflation targeting policy since the 1990s.

In this section, I argue that the move to unofficial inflation targeting since 1990 provides a natural setting to examine the seminal argument of Fogler (1984)—that the effectiveness of real estate as an inflation hedge is concentrated in high inflation environments. If this argument is valid, I hypothesise that the effect of inflation on housing markets will significantly reduce during low inflation periods or periods of inflation targeting. Though there is widespread consensus that residential real estate offers protection against realised inflation, little is known about whether the benefits and practice of inflation hedging vary according to the monetary environment. This section addresses this question.

6.2 Comparison of Inflation Hedging Across Sub-Periods

Table 11 reports the results of estimating Equation (7) over two sub-periods (1871 to 1989 and 1990 to 2012) to compare the inflation hedging capacity within and outside the inflation targeting regime. I find that inflation can account for 12.73% of the variation in housing price changes during the sub-period 1871 to 1889 but only 0.21% in the recent period of inflation targeting from 1990 to 2012. In contrast, real income explains 0.26% of variation in housing price changes in the non-inflation targeting regime but 26.13% in the more recent sub-period. My results suggest that the empirical relationship between macroeconomic fundamentals and housing markets is sensitive to the monetary environment.

Table 12 presents variance decomposition analysis based on the corresponding estimations in Table 11. In Column (4) of Table 12 with the three-factor setting from 1871 to 1989, inflation accounts for 87.92% of the explanatory power of 14.70%. However, in the three-factor setting with adjusted R-squared of 28.58% from the recent period of inflation targeting regime from 1990 to 2012, inflation only contributes to 2.02% of this explanatory power, while real income contributes to 93.46%. When country and time fixed effects are added in Column (7) of Table 12, the explanatory power of the model increases significantly in both sub-periods. This suggests that unobserved time invariant factors (i.e., country fixed effects) and systematic risk at the global level (i.e., time fixed effects) play a significant role in housing markets, irrespective of the monetary environments.

6.3 Implications

Though real estate is a physical asset with low liquidity and high transaction cost, it is commonly included in multi-asset class portfolios to diversify and protect against risk. It is well-known that real estate has not only a low correlation with financial assets, such as stocks and bonds, but also a high correlation with inflation. This paper examines inflation hedging characteristics in different monetary environments through a historical analysis of a century-long time series. In doing so, I show that the co-movement between inflation and housing returns is sensitive to monetary environments. Specifically, when the government attempts to stabilise prices using inflation targeting, it lessens the inflation risk for investors to hedge against. This does not suggest that real estate offers poor protection against inflation but rather highlights that the decision to hedge against inflation is a conscious decision depending on the perceived risk in the monetary environment.

Table 11. Panel data regression over sub-periods

	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Panel A: Sub-period from 1871 to 1989							
Constant	0.051***	0.038***	0.0473**	0.012***	-0.008	-0.002	0.002
	(0.004)	(0.004)	(0.006)	(0.006)	(0.014)	(0.054)	(0.052)
Income	0.130*			0.259***	0.192***	0.067	-0.012
	(0.067)			(0.063)	(0.062)	(0.073)	(0.072)
Inflation		0.362***		0.393***	0.345***	0.297***	0.230***
		(0.029)		(0.029)	(0.030)	(0.036)	(0.036)
Population			0.657	1.270***	2.114***	0.931*	1.913***
			(0.482)	(0.450)	(0.544)	(0.494)	(0.640)
Country fixed effects					✓		✓
Time fixed effects						✓	✓
Adjusted R-squared	0.26%	12.73%	0.08%	14.70%	18.91%	16.82%	21.56%
Observations	1,051	1,051	1,051	1,051	1,051	1,051	1,051
Panel B: Sub-period from 1990 to 2012							
Constant	0.016***	0.029***	0.024***	-0.004	-0.0065	-0.076***	-0.11***
	(0.004)	(0.007)	(0.008)	(0.008)	(0.0239)	(0.021)	(0.032)
Income	1.515***			1.564***	1.496***	1.688***	1.596***
	(0.167)			(0.165)	(0.165)	(0.224)	(0.224)
Inflation		0.347		0.364	-0.007	1.330***	0.896***
		(0.286)		(0.261)	(0.268)	(0.313)	(0.341)
Population			1.971*	2.054**	2.745*	1.692*	5.205***
			(1.075)	(0.987)	(1.619)	(0.946)	(1.568)
Country fixed effects					✓		✓
Time fixed effects						✓	✓
Adjusted R-squared	26.13%	0.21%	1.02%	28.58%	34.37%	42.85%	48.37%
Observations	230	230	230	230	230	230	230

Notes: This table presents the results of estimating Equation (7) over different sample periods. The sub-period from 1871 to 1989 is before the inflation targeting regime, whereas the sub-period from 1990 to 2012 is in the inflation targeting regime. ***, ** and * denote significance levels at 1%, 5% and 10%, respectively.

Table 12. Variance decomposition analysis over sub-periods

	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Panel A: Sub-period from 1871 to 1989							
Income	1	·	·	8.25%	4.22%	0.38%	0.01%
Inflation	·	1	·	87.82%	60.76%	31.48%	15.32%
Population	·	·	1	3.93%	6.74%	1.63%	3.33%
Country fixed effects	·	·	·	·	28.27%	·	24.28%
Time fixed effects	·	·	·	·	·	66.51%	57.06%
Adjusted R-squared	0.26%	12.73%	0.08%	14.70%	18.91%	16.82%	21.56%
Observations	1,051	1,051	1,051	1,051	1,051	1,051	1,051
Panel B: Sub-period from 1990 to 2012							
Income	1	·	·	93.46%	72.10%	36.29%	28.21%
Inflation	·	1	·	2.02%	0.00%	11.53%	3.82%
Population	·	·	1	4.52%	2.52%	2.04%	6.10%
Country fixed effects	·	·	·	·	25.37%	·	17.08%
Time fixed effects	·	·	·	·	·	50.14%	44.79%
Adjusted R-squared	26.13%	0.21%	1.02%	28.58%	34.37%	42.85%	48.37%
Observations	230	230	230	230	230	230	230

Notes: Based on the result in Table 11, I perform the corresponding variance decomposition analysis. I first compute the partial sum of squares for each effect in the model and then normalise each estimate by the sum across the effects, forcing each column for variance decomposition to sum to one.

7. Avenues for Further Research

Several extensions of my framework can be considered. First, my empirical results suggest that periods of inflation stabilisation contain little inflation risk for housing investors to hedge against. It would be interesting to explore whether such a source of inflation risk mainly comes from expected or unexpected inflation if a proxy for the expected inflation rate or a historical dataset of 30 day treasury bill rates is available.

Second, I focus on macroeconomic fundamentals in housing markets, which can be regarded as a broad extension of the macroeconomic factor model of equity markets in Chen, Roll and Ross (1986). While Chen et al. (1986) and this paper highlight the importance of inflation and productivity growth in understanding asset prices of equity and housing, respectively, a comparative study across asset classes would be of interest and importance to investors and policymakers.

8. Conclusion

The purpose of this paper is to investigate the inflation-hedging capacity of residential real estate in the long-run. I employ a novel dataset from Knoll et al. (2017) to model the dynamic relationship between macroeconomic fundamentals and house prices, focusing on inflation hedging characteristics under different monetary environments. While the seminal work of Fogler (1984) theoretically argues that real estate serves as a hedge most effectively in high inflation environments, we know little about the validity of this argument when applied empirically to real world data. Utilising the recent inflation targeting regime since 1990 as a natural setting, this empirical work seeks to better understand inflation protection from homeownership from 1871 to 2012.

In the context of three macroeconomic fundamentals (i.e., inflation, population and real income), I find that inflation is the only proposed factor that has a consistently positive and significant relationship with housing prices across all 10 countries examined. A further decomposition analysis of the three macroeconomic fundamentals in housing markets show that at least half of the explained variation can be attributed to inflation in 9 out of the 10 countries. This suggests that inflation is the best candidate among the known fundamentals to explain historical variation in housing price changes.

However, since the 1990s, I show that the dominant role of inflation hedging in housing disappears, which coincides with a time when most countries have achieved relative price stability through an inflation targeting policy. Inflation alone can account for 12.22% of variation in a panel analysis of 10 countries over the 1871 to 2012 period, but this significance reduces to 0.21% in the recent period from 1990 to 2012. The conventional view that housing serves as an inflation hedge empirically holds from 1871 to 2012 across 10 countries. However, this paper shows that hedging effectiveness is less likely to be observed in a regime of inflation targeting.

Overall, my results suggest that housing is an important hedging asset against inflation based on the international evidence from 1871 to 2012. My analyses across inflation and non-inflation targeting regimes indicate that if inflation does not take off, housing value would not increase and create the hedge. The empirical results contained in this paper warrant attention, since it indicates a potential monetary linkage between housing prices and inflation. The findings in this paper have broad implications for empirical asset pricing and policy research. From the perspective of empirical research, the role of inflation hedging is dynamic over time and thus the observed results can be sensitive to difference in monetary environments. From a policy point of view, my results imply that when policy makers stabilise inflation, little inflation risk would prevail in the market for investors to hedge.

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